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Explaining the Increase in On-the-Job Search

by Mikal Skuterud

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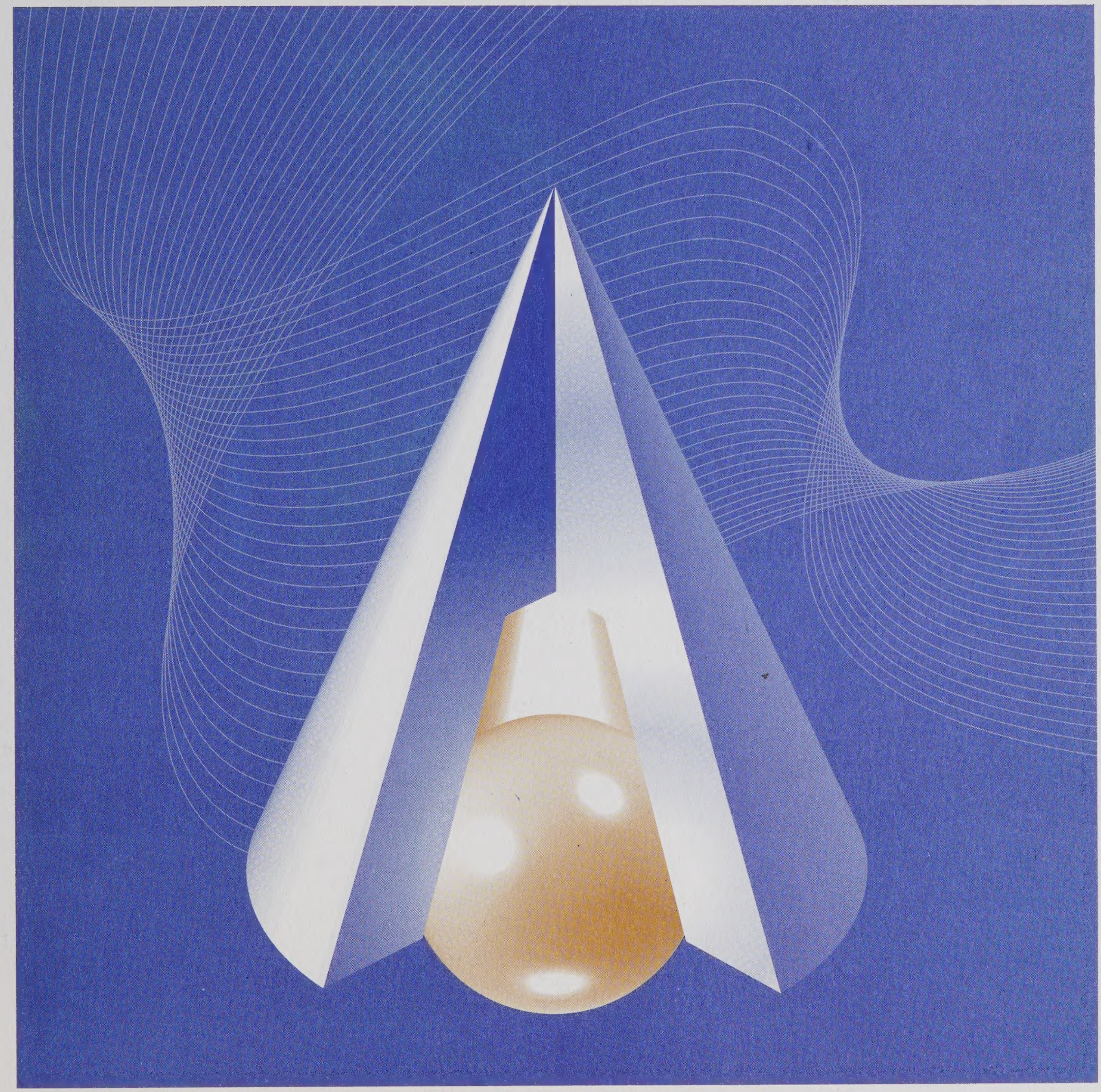

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Canada owes the success of its statistical system to a long-standing partnership between Statistics Canada, the citizens of Canada, its businesses, governments and other institutions. Accurate and timely statistical information could not be produced without their continued cooperation and goodwill.

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Abstract

Evidence from the Labour Force Survey (LFS) reveals that the percentage of employed workers searching for other jobs more than doubled in Canada between 1976 and 1995. Comparable evidence from the Current Population Survey (CPS), Panel Study of Income Dynamics (PSID), and National Longitudinal Survey (NLS) suggests that the U.S. experienced a remarkably similar upward trend in on-the-job search (OJS) over this period. Using U.S. data to supplement the Canadian data wherever possible, this paper attempts to explain this long-term, secular trend in Canadian OJS rates by performing decomposition and industry-level analyses, and by considering concomitant changes in employer-to-employer transition rates and the wage returns to job changing. The results from both countries suggest that an important part of the upward trend in OJS rates is not explained by compositional effects, including cohort effects. The OJS increase seems also to have occurred independently of rising job insecurity due to sector-specific demand shocks and trends in the dispersion of log wage residuals. The data are most consistent with a long-term decrease in search costs.

Keywords: labour turnover; wage differential

JEL: J630 J310

1. Introduction

There is a popular perception that today's workers are less inhibited by a sense of company commitment or loyalty than were their counterparts of the 1960s and 1970s. In the language of Albert Hirschman, the belief is that workers are increasingly likely to respond to job dissatisfaction with exit instead of voice or loyalty.[1] This trend has been attributed to at least two economic developments. First, workers perceive that the culture of lifetime jobs and joint employer-employee commitment was sacrificed during the recessions and widely reported corporate downsizing of the early 1980s and 1990s. Second, increasing education levels suggest that the relative importance of general human capital, as opposed to firm specific skills, may have grown significantly over the past three decades. In both cases, today's workers in their twenties and thirties are seen as behaving more like "free agents" than "company men," identifying themselves by their portable skills instead of the companies they work for.[2]

Despite the pervasiveness of these perceptions in both the popular press and human resources literature, the economics literature has struggled to find evidence of long-term changes in workers' job attachment. This is true whether it has focused on changes in job tenure distributions or directly on the incidence of either voluntary or involuntary job separations. The evidence from analyses of the regular job tenure data available in the Labour Force Survey (LFS) between 1976 and 1995 (a relative advantage of the Canadian data since the monthly Current Population Survey (CPS) from the U.S. does not collect tenure data) does not suggest a general increase in job instability for workers across the tenure distribution. Instead, there appears to have been a hollowing out of the middle of the tenure distribution with increased probabilities of both relatively short and long-term jobs (Green and Riddell 1996, Heisz 1996). Similarly, in a recent edited volume examining changes in the employment relationship in the U.S. during the 1990s, Neumark (2000) concludes that although there is some evidence of weakened employer-employee bonds, the magnitude of these changes has not been large enough to infer long-term trends.

This paper provides new evidence on a long-term change in employee loyalty to firms by focusing on what is perhaps a more robust indicator of a trend—the percentage of employed workers looking for other jobs. Using data from the LFS and three U.S. data sources with information on the search activities of employed workers—the Current Population Survey (CPS), Panel Study of Income Dynamics (PSID), and two cohorts of the National Longitudinal Survey (NLS)—the study constructs on-the-job search (OJS) rates for wage and salary workers between 1976 and 1995. The resulting data from both countries suggest that employees were slightly more than twice as likely to be searching for other jobs in the mid-1990s than they were in the mid-1970s. Decomposition analyses suggest that an important part of the trend in both countries cannot be explained by compositional shifts, including cohort effects. Rather, the increases in OJS appear to be true period effects experienced by all types of workers. These period effects seem to have occurred independently of rising job insecurity due to sector-specific demand shocks and concomitant increases in the dispersion of log wage residuals. From consideration of changes in employer-to-employer transition rates and the resulting wage returns

1. See Albert Hirschman (1971), *Exit, Voice and Loyalty*, Cambridge, Mass.: Harvard University Press.

2. This perception was recently expressed in *The Economist*'s "Survey of Youth", December 23, 2000.

to job changing over the period, the data appears most consistent with a long-term decrease in search costs.

There now exist three separate economic literatures concerned with OJS. The first is a surprisingly small empirical literature on the incidence and determinants of OJS. In the U.S. this includes a study by Rosenfeld (1977) using a special supplement to the CPS in May 1976, Black (1981) using a sample of household heads drawn from the 1972 PSID, and Meisenheimer and Ilg (2000) who use the contingent worker supplements to the CPS asked in February 1995, 1997 and 1999. For the U.K. it includes Pissarides and Wadsworth (1994), who use the 1984 British Labour Force Survey. The Canadian LFS data on OJS examined in this paper have not been examined elsewhere.

There is also an empirical literature concerned with the relative effectiveness of employed and unemployed job search. Blau and Robins (1990) and Holzer (1987) estimate reduced-form models of job search. An important difficulty in these studies is determining how much of the estimated differences are driven by individual unobserved heterogeneity in search abilities. In an attempt to correct for this heterogeneity, Jones and Kuhn (1996) focus on a sample of workers receiving different amounts of advance notice of layoff. Similarly, Burgess and Low (1992) examine whether preunemployment search intensity is increased by advance notice of layoff and Gottschalk and Maloney (1985) consider whether the choice of employed and unemployed job search affects job-finding probabilities for workers receiving advance notice. Finally, Belzil (1996) estimates a representative agent model of job search.

The third literature is the theoretical literature that has tried to incorporate OJS into models of unemployed job search. The challenge in this area has been to generate models with endogenous search strategies and both employed and unemployed job search in equilibrium. If one of these search strategies dominates in generating more wage offers or more favourable offers, such an equilibrium will obviously not exist. The original advances in this literature are the partial equilibrium models by Burdett (1978), Jovanovic (1979) and Mortensen (1986) that assume not all firms offer the same wage in any given period and both employed and unemployed workers are able to draw wages from the same nondegenerate wage distribution that describes these offers. The problem with this approach is that it ignores the demand side of labour markets. Burdett and Mortensen (1980) and Pissarides (1994) offer two quite different approaches to modeling search equilibrium with OJS.

The remainder of the paper is organized as follows. Section 2 discusses the data and trends of OJS in Canada and United States. Section 3 examines the role of compositional effects, including higher average educational achievement, increased incidence of contingent jobs, and cohort effects. In sections 4 and 5 industry-level analyses are used to explore the importance of sectoral reallocation trends and rising residual wage inequality respectively. Finally, in order to reconcile the results with the existing literature on trends in job instability and to provide some evidence on the relative importance of declining search costs, Section 6 considers concomitant trends in job-to-job transition rates and the resulting returns to job changing. The results are summarized in Section 7.

2. Data

The data used to construct OJS time-series for Canada and the U.S. are presented in Table 1. Evidence of an upward trend in the U.S. comes from five different series. First, in every year from 1969 to 1975 the PSID asked all employed household heads: "Have you been thinking about getting a new job, or will you keep the job you have now?" This question was dropped from the 1976, 1977 and 1978 surveys, but was then reintroduced in 1979 and kept until 1987 with *identical* wording and placement in the question ordering of the employment section of the PSID survey as it had been in 1975. There is therefore no reason to believe that the 1969-1975 and 1979-1987 are not comparable. In addition, in every year from 1979 to 1987 the PSID asked all respondents that were thinking about getting a new job the follow-up question: "Have you been doing anything in particular about it?" Responses to these two questions provide information on trends in OJS through the 1970s and first half of the 1980s.

In 1988 two important changes were made that affected the OJS information contained in the PSID. First, the two-part question was dropped in favour of the single question with a reference period: "Have you been looking for another job during the past 4 weeks?" Second, the sample was extended to include the wives of household heads. Although the latter change has greatly enhanced the value of the OJS data in the PSID by substantially raising sample sizes, the former served to break the continuous 1969-1987 series. Fortunately, in both May 1976 and May 1977 the CPS contained special supplements that asked all employed respondents with at least 4 weeks job tenure an essentially *identical* question to the new PSID question: "During the past 4 weeks, have you looked for another job?" To the extent that the PSID sample has maintained its national representativeness since its creation in 1968, a common sample of household heads and wives can be taken from the CPS to produce comparable OJS rates for 1976-1977 and post-1987. Using the CPS estimates as benchmarks, Fitzgerald, Gottschalk and Moffitt (1998) find that despite a 50 percent sample loss from the original 1968 PSID sample, there is no strong evidence that sample attrition has distorted the representativeness of the original PSID sample through 1989. The CPS and PSID data should therefore provide some additional evidence on trends in OJS over the past two decades.

Selected annual surveys from two cohorts of the NLS also contain OJS questions with a four-week reference period. The 1984 National Longitudinal Survey of Youth (NLSY) asked all employed workers: "Have you been looking for other work in the last 4 weeks?" Similarly, the 1985 and 1987 Surveys of Work Experience of Young Women (NLSYW) asked employed workers: "Have you been looking for other work during the past 4 weeks?" Again, comparable samples are taken from the 1976-1977 CPS and compared to the OJS rates from these NLS samples from the mid-1980s.[3] However, unlike the PSID comparison, we should expect the NLS rates to exceed those based on the CPS question, all other things unchanged, since the former questions refer to "other work" while the latter ask about "another job." Arguably, workers looking for *additional*, as opposed to *new*, jobs are more likely to respond affirmatively to the former. Due to this contrast in question wording the OJS data from the NLS are not used in the subsequent analysis.

3. The NLSY actually included the OJS question in every year between 1979 and 1984. However, this is a longitudinal data set of a particular cohort born between 1957 and 1965, so changes in the incidence of OJS over these six samples will include age effects in addition to any time-series trends.

The resulting OJS rates for the U.S. are plotted in Figure 1. The entire PSID and NLSY samples are used, including the low-income supplemental samples, and the CPS, PSID, and NLS weights throughout. However, the military supplemental sample is excluded from the NLSY since there is no comparable sample in the CPS. All rates are based on the sample of employed workers aged 16 and over excluding workers on temporary layoff. The search decision problem is likely to be very different for the self-employed, so the analysis also limits the sample from the outset to wage and salary workers. Finally, where the question specifies a four-week reference period, the sample is restricted to employees with at least four weeks of job tenure to insure the question is not capturing unemployed search activity. Despite the use of these substantially different data sources, all the series suggest the incidence of OJS roughly doubled between the mid-1970s and mid-1980s. The percentage "thinking about getting a new job" increased from about 12 in 1975 to 24 in 1985, while the percentage "looking for another job" went from about 4 to 8 between 1976 and 1988. As expected, responses to the "looked for other work" questions in the NLS seem to overstate the OJS increase found using the original two-part PSID question that refers specifically to search for new jobs. Taken together Panels 1 and 2 provide assurance that the increase in OJS reflects search for replacement jobs that involved actual search effort. It is also worth noting that the upward trend appears to predate the culture of downsizing that is typically linked to the recessions of the early 1980s and 1990s. This raises doubt about the role of corporate responses to these economic downturns in undermining company loyalty.

While the above evidence is clearly suggestive, these U.S. series are imperfect in terms of their breaks and the relatively small PSID and NLS samples. The Canadian series presented in Figure 2 is a substantial improvement over the U.S. in both respects. The LFS is a monthly nationally representative survey of the Canadian population providing sample sizes that exceed 25,000 in every month between 1976 and 1995 (mean sample sizes for all series are shown in Table 1). Further, the basic monthly LFS contained a question on OJS that was asked in every month of *all* employed respondents with positive hours of work during the reference week and remained unchanged throughout this twenty-year period. This study is the first to explore these data.[4] Using the March files of the LFS and the common sample of employed wage and salary workers, aged 16 and over, with at least 1 month tenure, these data reveal a remarkably persistent secular upward trend in Canadian OJS. Moreover, the Canadian series also suggest OJS rates approximately doubled between the mid-1970s and mid-1980s. Specifically, the rate increased from just over 2 percent in 1976 to slightly more than 4 percent in 1985. Although the rate continued to increase through the latter half of the 1980s and into the 1990s, there was a definite deceleration in its growth. So between 1985 and 1995 the rate changed by a single percentage point compared to the more than two percentage point increase in the previous decade. This also accords with the U.S. evidence using the considerably smaller sample size of the PSID. It is interesting that in all comparable years U.S. employees appear to be twice as likely to be looking for other jobs as Canadians. It turns out that decomposition analysis using information on respondents' industry, occupation, part-time status and various demographic characteristics is unable to explain any of this difference. This paper will focus on explaining trends and will leave the cross-country difference to future research.

4. Due to confidentiality concerns, the OJS variable was never included in the LFS public-use files between 1976 and 1995. Statistics Canada ceased to collect *any* information on the job search activity of employed workers with the LFS redesign in 1995.

Having established common upward trends in OJS in Canada and the U.S. between the mid-1970s and mid-1990s, it is worth considering how these increases in search activity varied between some demographic groups that can be consistently defined in the Canadian and U.S. data. Table 2 presents OJS rates and percentage point changes in these rates by gender/age group.[5] Consistent with the findings of Pissarides and Wadsworth (1994) and Meisenheimer and Ilg (2000), there is very little difference in search rates between men and women. There are some large differences between young men and young women in the U.S., but these disappear if the sample is not restricted to the PSID sample of household heads and wives. All three data sources also show higher OJS rates among younger workers. This is also consistent with other research and is typically explained in one of two ways: (i) younger workers have more years to collect the benefits of job changes or (ii) the quality of job matches tends to rise over the life-cycle as workers refine their skills and become more effective in collecting and evaluating job offers. On the other hand, the differences may simply reflect cohort effects as workers born in later years have more portable skills or a different sense of company loyalty. Table 2 also reveals that increases in OJS occurred across all age groups for both men and women. Measured as percentage point changes the increases tend to decline with age, while the percentage changes of the rates are rising with age. Unfortunately, with the exception of women aged 40-49 these changes are not statistically significant in the U.S. data. This reflects the small size of the 1993 PSID sample, rather than differences in the magnitude of the changes between the two countries. When the relatively large samples of the LFS are used all the changes are statistically significant at the 1% level. These results suggest that changes in the age distribution of the labour force cannot explain the trends, although they are of course entirely consistent with other compositional shifts, including cohort effects.

3. Compositional effects

3.1. Decomposition analysis

The simplest way to examine the role of other compositional shifts is to construct counterfactual changes in OJS rates by holding the distribution of observable characteristics constant in either the mid-1970s or mid-1990s.[6] To the extent that the size of the estimated increases in OJS rates between two years remain similar to the actual increases in Figures 1 and 2, the role of compositional shifts can be ruled out. Differences between three different pairs of OJS rates are decomposed using this technique: (i) the 1975 and 1985 U.S. rates taken from panel 1 of Figure 1, (ii) the 1976 and 1993 U.S. rates taken from panel 3 of Figure 1, and (iii) the 1976 and 1995 Canadian rates taken from Figure 2. Since each data source provides different demographic and labour market information, performing the decomposition in three ways improves our ability to explain the trends.

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5. The 1993, instead of 1995, PSID is used because the former is the most recent Final Release (Public Release II) file available. A number of variables used in the analysis here, such as industry codes, are not available in the Early Release files.

6. This general technique for decomposing predicted changes has been credited to Oaxaca (1973) and Blinder (1973), although they never used it in the context of a binary dependent variable model. For examples of decompositions in a probit model context see Even and MacPherson (1993) and Doiron and Riddell (1993).

The first stage in performing these decompositions involves estimating probit models that predict either the early or late year OJS rate separately for each of the three series. The results from using the early year cross-sections are presented in Table 3. Data from both countries show younger, more educated, part-time employees are significantly more likely to be searching for another job.[7] In addition, both the CPS and LFS suggest married women are significantly less likely to report OJS than single men. Some other noteworthy results are obtained from variables unique to one of the data sources. The PSID contains a series of questions on annual hours preferences. The results in column 1 reveal that employees with preferences for additional hours are significantly more likely to be looking for a new job than workers that are content with their current hours. This is consistent with the Altonji and Paxson (1988) finding that underemployed workers are relatively more likely to quit their jobs unless compensated for their short hours. Using information on reason for part-time employment the Canadian data also implies much higher probabilities of OJS for employees that feel under-worked. However, the PSID data suggest that demographically comparable workers that feel overworked are not significantly more likely to be searching for a new job. Perhaps surprisingly, workers paid on an hourly basis are significantly more likely to be searching, while union members appear, if anything, only marginally less likely to search. The former result probably reflects, at least in part, the strong tenure effects reported in column 3 that are not available in the U.S. data. These tenure effects are a well established result that Pissarides and Wadsworth (1994) attribute to low tenure workers having not yet ascertained the non-pecuniary characteristics of their jobs and having invested less in job-specific skills.

Having estimated probit models using each of the early year cross-sections, it is possible to predict OJS probabilities for each observation in these samples. The average predicted probability of OJS for year t, $\overline{p}_t$, is then simply:

$$\overline{p}_t = \frac{1}{N_t} \sum_i \Phi(X_{it} \beta_t) \qquad (1)$$

where N_t is the year t sample size, Φ is the normal cumulative density function, X_{it} is a vector of independent variables (all dummy variables) for observation i in year t, and β_t is the estimated coefficient vector from the probit model using the year t data. Since $\overline{p}_t$ equals the actual OJS rate in year t, the actual increase in OJS rates is $(\overline{p}_l - \overline{p}_e)$ where l and e index the late and early year respectively. This change can be decomposed into a part due to changes in characteristics, X_{it}, and a part due to changes in the returns to those characteristics, β_t. This is done by predicting the counterfactual rate:

$$\overline{p}_l^0 = \frac{1}{N_l} \sum_i \Phi(X_{il} \beta_e) \qquad (2)$$

7. Note the CPS defines part-time work as usual weekly hours less than 35, as opposed to less than 30 in the LFS. Analysis suggests that those working 30-35 hours have OJS probabilities more like the <30 hours group than >35 hours group. Except where the involuntary/voluntary distinction is made, the U.S. definition is therefore used.

and decomposing the increase in rates between the early and later year:

$$\bar{p}_l - \bar{p}_e = (\bar{p}_l - \bar{p}_l^0) + (\bar{p}_l^0 - \bar{p}_e) \tag{3}$$

where the first and second terms are usually referred to as the unexplained and explained parts of the difference respectively. Since it is equally legitimate to base this decomposition on probit estimates using the later year cross-sections, there are of course two ways of doing this analysis.[8] The results in part A of Table 4 present estimates from both approaches. The results from the three data sources are remarkably similar. The upper bound U.S. and Canadian estimates using the mid-1970s and mid-1990s data are in fact identical. In each column the estimates suggest that compositional shifts can, at most, explain one-third of the increase in OJS experienced in the two countries.

It is possible to further decompose the explained part of the trend into the change due to each of the independent variables. This is done by allowing only the elements of β associated with a particular variable to change when calculating the counterfactual rate. The complication here is that in a nonlinear model, like the probit, the calculated changes due to each variable will not, in general, add-up to the total explained change. In order for the effects of the individual variables to add-up, it is necessary to linearize the probit function in some way. Following Even and MacPherson (1993), the approach used here is to attribute

$$(\bar{p}_l^0 - \bar{p}_e)_j = (\bar{p}_l^0 - \bar{p}_e) \cdot \frac{(\bar{X}_{jl} - \bar{X}_{je})\beta_{je}}{(\bar{X}_l - \bar{X}_e)\beta_e} \tag{4}$$

to the variable j set of dummy variables. The first term on the RHS is the total explained change, while the second term is the proportion of the total index change due to the variable j.[9]

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8. With the Canadian data it is also possible to use years other than 1976 and 1995. Although not reported, separate decompositions of the 1976 rate and every year between 1977 and 1995 were performed. The results were remarkably similar between specifications, emphasizing that there is little to gain from focusing on shorter periods within the twenty-year series.

9. Doiron and Riddell (1994) point out that a weakness of this technique is that it ignores the nonlinearty of the probit function. They argue that a preferred approach is to predict at the means of X_{it} and take a first order Taylor series approximation of the probit function. The explained gap can then be approximated by:

$$\Phi(\bar{X}_l \beta_e) - \Phi(\bar{X}_e \beta_e) \approx \frac{\partial \Phi(\psi)}{\partial \psi} \cdot (\bar{X}_l - \bar{X}_e) \cdot \beta_e \tag{5}$$

where the first term on the RHS is the derivative of the probit function evaluated at an arbitrary location ψ. And the contribution of variable j to this change is given by:

$$\frac{\partial \Phi(\psi)}{\partial \psi} \cdot (\bar{X}_{jl} - \bar{X}_{je}) \cdot \beta_{je} \, . \tag{6}$$

The results from using the late year parameter estimates are presented in part B of Table 4. Again the three data sources provide remarkably similar results. In all cases, the most important compositional change was the shift towards a more educated labour force. In all years, employees with post-high school education are significantly more likely to be searching and between the mid-1970s and mid-1990s their share increased from about 35 to 51 percent in the U.S. and 34 to 48 percent in Canada.[10] This provides some support for the hypothesis that the acquisition of more portable skills has resulted in weakened company loyalty. However, the data in columns 2 and 3 suggest that this shift can account for no more than a sixth and a third of the U.S. and Canadian trend respectively. In addition to educational changes, shifts to part-time, non-unionized, service sector jobs appear to have played a modest role in raising OJS.[11] Perhaps surprisingly, shifts in the tenure distribution appear to have reduced, not increased, search activity among employed workers. The explanation is that there was a bifurcation of the tenure distribution but, at least when the tenure distribution is truncated at 1 month, the increase in the probability of being in a long job dominates the increased probability of being in a short job. Finally, both U.S. data sources imply that shifts in the age distribution and marriage rate served to increase OJS rates. These results are a consequence of the sample restriction to household heads and household heads and wives. By excluding the single dependents, whom have low OJS probabilities relative to single heads, the estimated marital effect becomes negative and significant. This combined with an increase in the single share gives us the estimated effects. The unexpected age effect is more complicated but reflects estimated age effects that decline less smoothly than when the sample is not restricted to household heads and wives.

3.2. Contingent jobs

If shifts to a more educated labour force and more part-time, non-unionized, service sector jobs can explain only a third of the upward trends in OJS, then what explains the residual trend? A weakness of the decomposition analysis above is that it cannot account for compositional shifts that are unobserved or OJS effects that are not measured. Of particular concern is the absence of information on contingent work status in the data. There is a widespread belief and some limited evidence that the incidence of temporary and contract jobs has increased significantly over the past two decades (see Polivka (1996) and Segal and Sullivan (1997)). Since workers in

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Doiron and Riddell (1994), and more recently Morrisette and Drolet (2001), present results from both techniques. However, it is unclear that there is a substantive difference. Expressed as percentage point changes, the two approaches will produce different results, but expressed as the percentage of the total explained change due to each variable *j*, they produce *identical* estimates. The reason is that the derivative of the probit function is a constant so that (6) divided by the RHS of (5) produces exactly the second term on the RHS of (4). An alternative and simpler approach, taken by Morrisette and Drolet (2001) and Hamermesh (2001) that produces different results is to linearize by estimating a linear probability model. The results in part B of Table 4 are not substantively different from those obtained from a linear probability model.

10. In January 1989 the LFS coding of the education variable changed from one based on years completed to highest level attained. The consequences of this change are discussed extensively in Bar-Or et al. (1995) and their approach to coding is followed.

11. Shifts to part-time work are considerably more important in explaining the Canadian data. The reason is simply that the U.S. has not experienced a comparable increase in part-time rates over this period. Similarly, decreasing union density rates is important in explaining the U.S. increase in OJS, but not the Canadian. Again, the reason is simply that Canada did not experience a similar reduction in union density over this period.

contingent jobs are more likely to anticipate their terminations, we expect them to exhibit significantly higher OJS probabilities than similar workers in less flexible work arrangements. Since contingent worker rates may be only weakly correlated with the industry, occupation, weekly hours and demographic variables observed, the Section 3 estimation will attribute too little to compositional shifts. We do however expect the incidence of contingent work to be highly correlated with the tenure data available in the LFS. This is confirmed by estimates from the Survey of Work Arrangements, a special supplement to the LFS in December 1995. Among employees with less than one year of job tenure, slightly more than 30 percent claimed to be in a job that was in some respect not permanent. The comparable rate for workers with more than ten years tenure is less than 2 percent. Changes in OJS rates among high tenure workers should therefore shed some light on the importance of shifts to contingent work.

Table 5 presents OJS rates for a group of workers we expect to have strong job attachment—high tenure men with usual weekly hours of 40 or more. In order to boost the annual sample sizes, the twenty March LFS files between 1976 and 1995 are pooled into four periods. The estimates reveal increasing OJS across the tenure distribution. Although the percentage point changes are monotonically decreasing in tenure, measured as proportions of the first-period rates these changes are considerably greater among the high tenure employees. Assuming a simple binomial distribution with variance given by $p(1-p)/N$, these increases are all highly statistically significant. It is telling that even men with more than 10 years job tenure were more than twice as likely to be looking for a different job in the mid-1990s than in the mid-1970s. Clearly, an important part of the upward trend in OJS is not explained by shifts to contingent work. More generally, distributional shifts to other types of low-quality jobs with high turnover rates, whether explicitly fixed-term or not, are unable to explain the residual trend.

3.3. Cohort effects

It is still possible that no individual worker's commitment fell. Although there is only weak evidence of cohort effects based on workers' perceptions of corporate commitment to employees and higher educational attainment, they may still exist for different reasons. Perhaps workers born in the liberal political culture of the 1960s acquired a different sense of company loyalty than workers born before 1940. Table 6 presents OJS rates for four different cohorts in the United States. In order to raise sample sizes, the two CPS samples from the 1970s and the six PSID samples from 1988-1993 are pooled separately. The first feature to note from Table 6 is that the OJS rates are monotonically increasing in cohort. So even without any change in individual worker behaviour we should find increasing OJS rates when we compare cross-sections from the 1970s and 1990s. What happens to the trends when we look within cohorts? Consistent with a cohort effect explanation, the large increases in OJS seem to disappear. With the exception of the youngest cohort, all the period-to-period changes are statistically insignificant despite all the cell samples exceeding 3,500.[12] These results come across even more convincingly in the LFS data. Figure 3 plots the Canadian OJS series for the same four cohorts. Again, OJS rates are increasing in year of birth and the upward trends disappear when we look within cohorts. With the exception of a modest increase in the early period for the youngest cohort, the within-cohort profiles are remarkably flat relative to the total trend in Figure 2. On

12. Note that in the period 1976-1977 the youngest cohort group includes only 16 and 17 year-olds so these samples are relatively small in both the U.S. and Canadian data. This explains the different patterns for this cohort.

the surface, these results seem to suggest that cohort effects can account for all of the upward trends in OJS.

The problem is, of course, that these results follow directly from the two main results in Table 2—OJS rates are increasing over time and decreasing in age. With year and age effects of equal magnitude, within-cohort profiles must be flat. It is well known that since the year, age and cohort variables are linearly dependent (cohort = year – age), it is impossible to correctly identify all three without knowing something about the form of at least one of them. Fortunately, there is good reason to believe that independent negative age effects exist. In the context of the standard search model with OJS (see Burdett 1978 or Mortensen 1986), if employed workers are able to draw wages from a nondegenerate wage distribution in every period of their lifetimes, the probability of drawing a wage that exceeds their current wage, their best draw to date, must decrease as the number of draws or periods increase. The probability that optimizing agents will absorb the costs of search must therefore decline as the number of periods, or their age, increases. Further, where agents are finite-lived the expected benefits of OJS will fall simply because the present value of any wage improving draw declines with age. If we accept that independent negative wage effects exist, we should then expect the within-cohort profiles in Figure 3 to be declining. What explains the fact that these profiles tend to be flat? Clearly, there must have been period effects of roughly equal and opposite magnitude. Since we believe the age effects must be quite large, the period effects must similarly be important. This suggests that compositional shifts, including popular perceptions of cohort effects, cannot explain an important part of the upward trends in OJS rates. Rather, the residual trend appears to be true period effects, experienced by workers across the age and cohort distributions.

4. Sectoral reallocation

In an effort to determine the role of two well-established labour market trends that are plausibly related to the secular trend in OJS, in this section and the next, the study exploits between-industry variation in the growth of OJS. In this section the role of sectoral reallocation trends is examined with particular interest in the impact of the de-industrialization process experienced in Canada and the U.S. through the 1980s and 1990s. As argued by Lilien (1982) in his explanation of cyclical unemployment, shifts of employment demand between sectors of the economy necessitates job reallocation and therefore variation in the incidence of job search. Following this idea, if workers in declining industries anticipate their layoffs and workers employed in expanding industries realize new job opportunities, we might expect within-industry employment changes over a period to be correlated with the level of OJS. More formally, the probability that individual i employed in industry j searches for a new job in period t might be determined by:

$$\Pr(OJS_{ijt}) = f\left(X_i, \left|\Delta E_{jt}\right|\right) \tag{7}$$

where X_i is a vector of person-specific characteristics, E_{jt} is the employment level in industry j in period t, and $f_2 > 0$. This implies that:

$$\Delta \Pr(OJS_{ijt}) = g\left(\left|\Delta E_{j,t+1}\right| - \left|\Delta E_{j,t}\right|\right) \tag{8}$$

assuming the terms in (7) are additively separable. Evidence on the role of sectoral reallocation can then be obtained by considering whether there is a positive correlation between within-industry changes in OJS and within-industry differences-in-differences employment changes.

A complication in this analysis is knowing how far into the past workers look when making search decisions based on employment level changes. Ideally a number of different time horizons could be explored, but unfortunately the analysis is limited by the available OJS data. Using the LFS data, which offer the more complete OJS time-series, the study constructs absolute annual employment level changes at the two-digit industry level from March-to-March of every year between 1976 and 1995. To the extent that (7) explains OJS behaviour, these employment level changes should be correlated with OJS rates observed in the March LFS files of each year. A consequence of the level of industrial detail is that some of the cell sizes, particularly among the manufacturing industries, are worryingly small. In response the focus is on mean OJS rates and employment changes within four five-year periods: 1976-1980, 1981-1985, 1986-1990 and 1991-1995. As long as f_{22} is close to 0 over the range of employment changes in the data, this aggregation is innocuous. In addition, in all cases the OLS estimates of the correlations are supplemented with estimates from a weighted least squares (WLS) procedure, which serves to weight observations according to the size of the samples used to generate the industry-specific rates. The details of this estimation are given in Appendix B.

Before examining the correlations, it is worth comparing the changes in OJS experienced within Canadian industries over the four five-year periods. These rates for 45 two-digit industries are presented in Table A1. At least two noteworthy results emerge from these data. First, the OJS increases are remarkably pervasive, as no industry appears to have experienced a net decline in search activity between the late-1970s and early-1990s. Second, the increases are on average larger for service-producing industries than goods-producing industries. Given the shift away from manufacturing employment over this period, this casts some doubt on the role of heightened job insecurity in raising OJS.

Figure 4 presents the correlations between changes in OJS and differences-in-differences employment changes. Given the contrasting employment experiences of the manufacturing and service sectors over this period, the panels of this figure show this relation separately for goods and service-producing industries. Clearly the results do not support the hypothesis that sectors experiencing relatively large increases in employment instability should see the largest increases in OJS. This casts doubt on the importance of sectoral reallocation trends in explaining the rising OJS rates. However, perhaps the search decision in (7) is based on lagged values of OJS_{ijt} and longer-term employment changes. We should then see a u-shaped relationship between first-difference employment changes and increasing OJS rates. Figure 5 considers this correlation using both the Canadian and U.S. data. In order to make the Canadian results comparable to the U.S., the study focuses on the period from the mid-1970s to the late-1980s when aggregate OJS rates appear to have roughly doubled. The OJS rates for 37 two-digit industries in the U.S. are shown in Table A2. Similar to the Canadian within-industry OJS changes, these data reveal remarkably pervasive increases that are on average larger among service than goods-producing industries. When plotted against first-difference employment changes in Figure 5, there now appears to be some role for sector-specific demand shocks in explaining the OJS increases. Although the results for goods-producing industries are all statistically insignificant, the service-

producing estimates from both countries now suggest a marginally significant positive correlation. Arguably, since there are relatively few industries that experienced employment contractions over this period, it is not surprising that a negative correlation does not appear.[13]

However, closer scrutiny of these results reconfirms our doubt of the role of sectoral reallocation trends. First, despite the positive correlations in Figure 5, the estimated relationships imply rising OJS even where employment levels are stable. As Figure 6 reveals, much of the positive correlation that is observed can, at least for Canada, be explained by increasing part-time and contingent employment rates in the service sector. Second, despite experiencing widespread negative demand shocks through the recession of the early 1990s, data from the Canadian manufacturing sector over this period also does not suggest a role for rising job insecurity. The first three panels of Figure 7 consider changes in production worker employment levels using the Annual Survey of Manufactures (ASM). The annual employment loss rates, plotted in the third panel, are calculated for each industry as the loss in employment from all plant contractions and closings between two years as a function of the total employment in the first year.[14] Regardless of the measure used, both the OLS and WLS estimates are now statistically insignificant despite all the industries, except rubber and plastics, experiencing negative employment changes over this period. Further, changes in the mean production worker wage appear to be, if anything, *positively* correlated with OJS increases. Finally, the fifth and sixth panels of Figure 7 consider the level and changes in the ratio of imports to domestic consumption by industry.[15] Again, regardless of the measure used the estimates are highly insignificant. There is clearly no evidence that heightened job insecurity due to increasing import competition or any other negative demand shocks served to raise OJS. Taken together these results suggest that something other than sectoral reallocation trends is responsible for the observed change in search behaviour among U.S. and Canadian employees.

5. Firm wage effects

An alternative explanation of the period effects found in Section 3 comes from consideration of within-industry earnings inequality. Consider the log earnings equation:

$$\log w_{ijt} = X_{ijt}\beta_{jt} + u_{ijt} \qquad (9)$$

where w_{ijt} is the weekly earnings of individual i, in industry j, in year t, X_{ijt} is a set of observable characteristics, including human capital characteristics, and u_{ijt} is an iid random error term. In addition to weekly hours of work and the unobserved ability of individual i, u_{ijt} should capture

13. As plots in Figure 5 suggest, pooling the goods and service-producing industries and adding a quadratic term produces no evidence of a negative correlation over the range of negative log employment changes.

14. These data were generously provided by John Baldwin from the Micro-Economics Analysis Division at Statistics Canada.

15. These ratios are constructed using the import data by commodity in the System of National Accounts and the assumption that an industry's lost production due to trade is equivalent to its domestic production share of each imported commodity.

firm wage effects. These include noncompetitive factors, such as unionization, monopoly status in the labour market and discrimination, efficiency wages, and varying returns to unobserved ability between firms. Regardless of the source of these firm wage effects, it is conceivable that a change in their distribution within an industry results in increased OJS as workers seek wage improvements through within-industry job changes. As evidence that firm wage effects are an important determinant of OJS behaviour, Bhaskar et al. (2002) show that workers in higher-paid jobs are less likely to search for other jobs and search less intensively when they do than otherwise similar workers in lower-paid jobs. Increased dispersion of industry-level weekly hours could also serve to raise search activity, although this is likely to be less important given the focus on full-time workers. Yet, in either case OJS trends driven by firm wage effects should be captured by changes in the distribution of u_{ijt}. In contrast, changes in the distribution of unobserved ability, due to more unequal schooling quality for example, should not raise OJS, although it may affect the distribution of u_{ijt}. Although we cannot separately identify the firm and hours effects component of u_{ijt} from the unobserved ability component, evidence of a positive correlation between within-industry residual wage inequality and within-industry increases in OJS is certainly suggestive of firm wage effects.

Using the March files of the CPS from 1977-1989 and the Survey of Consumer Finance (SCF) individual files from 1981 to 1996, equation (5) is measured separately for 36 and 13 industries in the U.S. and Canada respectively.[16] Both of these surveys provide retrospective information on work activity and earnings in the previous calendar year, including total earnings from wages and salaries and total weeks worked. These data are used, along with CPI indexes from each country, to construct real weekly earnings.[17] The samples are restricted in all cases to full-time, full-year workers aged 16 to 65 with positive wage and salary earnings in the reference year. In an effort to avoid self-employment income as much as possible, the agricultural industry is dropped and the study limits the U.S. data to workers whose longest job in the reference year was a wage and salary job and the Canadian data to workers whose major source of income was wages and salaries.[18] The vector of observable characteristics includes information on employees' education, labour market experience, region/province, sex, marital status, urban/rural residence, and occupation.[19] In addition, in the U.S. it includes dummies for black and Hispanic and in Canada for immigrant status and French as a mother tongue. Finally, in order to avoid

16. Unfortunately, the most detailed industry code available in the Survey of Consumer Finances identifies only 13 industries.

17. The U.S. CPI series are annual, seasonally unadjusted, all items, base year 1982-1984. The Canadian series are annual, seasonally unadjusted, all items, base year 1992.

18. It turns out the results are quite sensitive to the inclusion of the agricultural industry. In particular, in both countries the agricultural industry experienced an unusually large reduction in the dispersion of firm wage effects (note that with the exception of the personal services industry in the U.S., the agricultural industry had the highest average level of firm wage effects dispersion in both countries over the periods examined). Coupled with very mild OJS increases in both countries, the inclusion of this industry suggests a weak positive correlation between changes in the dispersion of firm wage effects and increases in OJS. As Figure 9 reveals, the relationship is negative when agriculture is excluded. Given the difficulty of interpreting firm wage effects in this industry, the decision was made to exclude agriculture.

19. Three occupation dummies are constructed in both the U.S. and Canadian data: (i)white-collar (managerial-adminisitrative-professional), (ii) pink-collar (clerical-sales-service), and (iii) blue-collar.

complications that result from outliers and the topcoding of the earnings data in the CPS, equation (5) is estimated by median regression (i.e., least absolute deviations) and the 90th minus 10th percentile of the absolute residuals is used as a measure of within-industry dispersion of firm wage effects.[20]

The aggregate U.S. and Canadian trends in these measures of dispersion in firm wage effects are plotted in Figure 8. They indicate remarkably similar magnitudes and upward trends during the early 1980s. Further, the U.S. data reveal that the trend predates the recessions of the early 1980s, which is hopeful for attempts to explain the OJS trends in Figures 1 and 2. However, these trends appear to diverge in the mid-1980s as the Canadian residual inequality seems to fall during the second-half of the 1980s. The absence of an upward trend in Canada through the 1980s raises some doubt about the ability of firm wage effects to explain the data.

Before attempting to explain the trends in OJS, the two panels on the left hand-side of Figure 9 correlate the average within-industry level of OJS with the average level of within-industry residual wage dispersion over the period of interest. The U.S. and Canadian results are again remarkably similar, both suggesting an important role for firm wage effects in explaining OJS behaviour. Clearly, employees in industries with higher levels of residual wage inequality, such as personal services in both countries, are more likely to search for new jobs than employees in industries with lower levels of residual wage inequality, such as the durable manufacturing industries in each country. Since it appears that OJS decisions are, at least in part, motivated by firm wage effects, it seems reasonable to expect upward trends in residual wage inequality to explain the observed increase in OJS. The two right-hand panels of Figure 9 plot this correlation. Although the sign of the point estimates from the two countries contrast, all the estimates are statistically insignificant. It seems that increased OJS has occurred even where the dispersion of firm wage effects has been declining. This suggests that firm wage effects are also unable to account for the widespread increases in OJS documented in Section 3.

6. Employer-to-employer transitions

The finding that heightened OJS activity appears to have occurred independently of broader labour market trends that should have changed the relative benefits of changing jobs, raises the important question of whether increased OJS has actually resulted in more employer-to-employer transitions. It is of course possible that employees are becoming increasingly fastidious about the jobs they accept at the same time as they search more. It seems plausible that falling search costs could raise the incidence of search *and* reservation wages so that OJS rates increase with little or no change in employer-to-employer transition rates.[21] Indeed, as noted above, the

20. Alternative measures have been estimated using the SCF data, which has earnings data that are not topcoded. These include taking the variance of the residual from an OLS regression and taking the variance, mean and median of the absolute residual from a median regression. In all cases, the main results are unaffected by the choice of measure.

21. The reason that this theoretical prediction does not exist is that it requires a model with both a search decision (as in Burdett (1978)) and a job acceptance decision (as in Hey and McKenna (1979)). The theoretical OJS literature has focused on either search or mobility costs, so that a model with both decisions does not exist. In

large literature concerned with changes in job instability over the past two decades has struggled to find evidence of long-term trends. The problem is that, with only two exceptions, this literature has focused exclusively on tenure distributions and overall job separation rates.[22] Interestingly, both papers that estimate employer-to-employer transition rates do find evidence of increasing instability. First, using the PSID from 1981 to 1992, Gottschalk and Moffitt (2000) compute separation rates conditional on exit destination and find some evidence of increasing transitions to non-self-employment through the 1980s. Second, using the March CPS from 1976 through 2001, Stewart (2002) identifies a dramatic long-term increase in employment-to-employment transition rates (job changes with two or fewer weeks of intervening unemployment) of about 50 percent. Interestingly, both papers also find similar increases for men and women and an offsetting decline in transitions to non-employment. The former result is consistent with OJS trends, while the latter reconciles these results with the extensive evidence of long-term stability in overall tenure distributions and separation rates.

Evidence of an upward trend in employer-to-employer transition rates and information on how the wage returns to job switching have changed provides useful insights into the cause of increased OJS. Unfortunately, neither the LFS in Canada nor the PSID and CPS in the U.S. can be used to construct these data between the mid-1970s and late-1980s when OJS rates roughly doubled.[23] Following Monks and Pizer (1998) and Bernhardt et al. (2000), who focus on two-year total job separation rates, the study uses two separate cohorts of the NLS to compare one-year transition rates by destination between the mid-1970s and late-1980s. This is possible because the National Longitudinal Survey of Young Men (NLSYM) and the National Longitudinal Survey of Youth (NLSY) cohorts were aged 23 to 31 in 1975 and 1988 respectively. Besides providing information on wages in both the old and new jobs, the NLS are a preferred data source because they contain unique employer identification codes, which are used to identify transitions. Brown and Light (1992) find that these employer codes are the best source of employer identification both within the NLS and when compared to other longitudinal data sets.

Figure 10 presents one-year transition rates using the sample of employed men from each cohort that are wage and salary workers in their main jobs. The poor white and military supplemental samples were excluded from the NLSY cohort, since there are no comparable supplemental samples available in the NLSYM. The weights from both cohorts are used throughout so as to produce representative samples of the 23-31 age group in each of the two years. Consistent with the results based on the PSID and CPS, the bottom-left panel of Figure 10 reveals an increase of almost 50 percent in the probability of making a job change despite a relatively small increase in

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the absence of search costs optimizing agents will always choose to search and in the absence of mobility costs the reservation wage is always equal to the current wage and therefore independent of search costs.

22. See Table 1.1 in Neumark (2000) for a useful summary of the measures and findings of this literature.

23. The LFS contained no wage information between 1976 and 1995. The PSID contains wage data, but depends on tenure data to identify job changes. The tenure questions have changed over time making rates from the 1980s incomparable to those from the 1970s (see Polsky (1999) for a summary of these changes). The March CPS, on the other hand, depends on retrospective questions about job changes in the past year. Wages on the old job are not observed and would probably be subject to tremendous measurement error if respondents were asked to recall these wages.

the overall job separation rate (shown in the top-left panel). Again, the contrast is explained by an offsetting decline in the probability of making a transition to non-employment (top-right panel). Interestingly, there is no indication that the increase in employer-to-employer transitions was disproportionately voluntary (bottom-right panel).

Given the potentially important compositional differences between the NLSYM and NLSY samples of young adult men, it seems important to adjust for these differences before implying a behavioural change. Table 7 presents the results from predicting the probability of experiencing an employer-to-employer transition by probit conditional on being observed in 1988, instead of 1975, and a set of characteristics whose means may have changed between these years. Regardless of the set of controls used a statistically significant increase in the incidence of making an employer change is found. As some assurance of the meaningfulness of these transitions, the estimates also suggest higher rates of employer change among younger, single workers and the wider the window between interviews. Together with the CPS and PSID results, there is strong evidence that rising OJS has resulted in more employer-to-employer transitions. However, the increase in OJS does not appear to be equivalent to the increase in job changing. This is entirely consistent with an increase in reservation wages motivated by a long-term decline in search costs.

Although the evidence in Section 5 suggests that the OJS increase was not caused by increasing dispersion of firm wage effects, it is still possible that workers have seen growth in the wage returns to job changing. The explanation based on falling search costs and rising reservation wages implies this result. Figure 11 presents kernel density estimates of the distribution of one-year real log wage changes experienced by employer-stayers and switchers separately for the 1975 and 1988 samples. As expected there is noticeably less dispersion of wage changes among the samples of employer-stayers than switchers. The results also indicate a higher probability of experiencing a positive wage change among the voluntary, than involuntary, switchers. However, reconfirming the irrelevance of wages in Section 5, there is no indication of a change in either the distribution of wage returns to job changing itself or its location. Consistent with the evidence on firm wage effects, the OJS increase does not appear to have been driven by workers anticipating improved wage returns to job switching. The results are also at odds with the explanation based on falling search costs and rising reservation wages. Yet, it is still entirely possible that workers have become increasingly fastidious in response to falling search costs. As emphasized by Blau (1991) in his empirical rejection of the reservation wage property, the search decision is more accurately based on a reservation utility decision. Surely workers also value the nonwage attributes of jobs such as hours of work (Blau's focus), fringe benefits, and working conditions. This suggests reservation utilities might have been rising in response to declining search costs. Since workers' utilities are not directly observed, obtaining evidence of this behavioural change is not straightforward.

Additional support for the hypothesis of declining search costs and rising reservation utilities becomes apparent when we recognize that the OJS rates in Figures 1 and 2 are based on stock samples of employed workers. This means that the computed rates will be sensitive to increases in both the incidence *and* duration of OJS spells. The fact that employer-to-employer transition rates appear to have increased implies that at least part of the OJS trend reflects a higher incidence of OJS spells. Longer OJS spells resulting from lower search costs and rising

reservation utilities offers an explanation for why stock sample OJS rates increased more than transition rates. In contrast, explanations based on declining mobility costs, such as more portable pensions or higher marital separation rates, imply lower reservation wages and therefore shorter durations and *lower* stock sample OJS rates (see Hey and McKenna (1979) for this prediction). Improved search efficiency, modeled as an increase in the offer arrival rate, due to improved communication technology or the introduction of private employment agencies should similarly reduce search durations and therefore stock sample OJS rates. Given these contrasting effects on stock sample OJS rates, the evidence appears most consistent with a long-term decline in the cost of OJS.

Unfortunately, it is difficult to think of explanations of why search costs have fallen in such a smooth way over the period from the mid-1970s to the mid-1990s. Since there is no foregone income cost of OJS, the most obvious explanation is that the use of a particular search technology has become more affordable. Between 1976 and 1981, the LFS also asked respondents reporting OJS to report their method of search. Table 8 considers changes in OJS rates by method over these six years. Although most of the OJS trend reflects increases in direct employer contact, the rates suggest that the use of public employment agencies, ads and other methods for OJS were also becoming more common. It is hard to believe that scanning newspaper ads became gradually less costly in the second-half of the 1970s. This suggests that something other than a change in the cost of using particular search technologies has been driving the OJS increase. Although there is no direct evidence of weakening company loyalty, the data are entirely consistent with decreasing OJS costs associated with foregoing loyalty to a current employer. These costs may be psychic, but need not be. Similar to the cost of shirking in the Shapiro and Stiglitz (1984) efficiency wage model, perhaps workers face a positive probability of being caught searching, which may be punished through lower future wage gains, loss of promotion possibilities or in the extreme case through job termination. The upward trend in OJS then implies that either the probability or the cost of being caught searching have gradually declined since the mid-1970s.

7. Summary

This paper is the first to examine long-term trends in the rate of on the job search in Canada and the United States. The nationally representative data examined reveal that the percentage of employed workers looking for other jobs more than doubled in both countries between the mid-1970s and mid-1990s. Analysis suggests that an important part of these trends cannot be explained by compositional effects, including cohort effects. Rather, the upward trends appear to reflect true period effects, in the sense that increased search activity occurred among all types of employees. This result is at odds with the popular perception that today's younger generations have a weaker sense of commitment or loyalty to their employers than their parents or grandparents did. Instead, their parents' loyalty seems to have fallen as much as their own.

To determine the cause of these period effects, variation in within-industry growth of OJS rates is correlated with variation in within-industry changes in employment, wages, import penetration ratios and earnings inequality. The results suggest that the period effects occurred independently of rising job insecurity due to sector-specific demand shocks and concomitant increases in the

dispersion of firm wage effects. From consideration of changes in employer-to-employer transition rates and the resulting wage returns to job changing over the period, the data appear most consistent with a long-term decrease in search costs. Although there is no direct evidence of the source of the decline in search costs, the data are entirely consistent with a reduction in the psychic costs of foregoing loyalty to a current employer or in the probability or cost of being caught searching while employed.

Table 1 **Data series**

Survey	Date	Sample (mean sample size)	Question
United States			
Panel Study of Income Dynamics (PSID)	1969-1975 and 1979-1987	Household heads that are wage and salary workers (3,023).	"Have you been thinking about getting a new job, or will you keep the job you have now?"
Panel Study of Income Dynamics (PSID)	1979-1987	Household heads that are wage and salary workers (3,153).	"Have you been doing anything in particular about it?"
Current Population Survey (CPS)	May 1976 and 1977	Wage and salary workers with tenure > 1 month (39,618).	"During the past 4 weeks, have you looked for another job?"
Panel Study of Income Dynamics (PSID)	1988-1995	Household heads and wives that are wage and salary workers with tenure > 1 month (3,999).	"Have you been looking for another job during the past 4 weeks?"
National Longitudinal Survey of Youth – (NLSY)	1984	Wage and salary workers aged 19 to 26 with tenure > 1 month (5,972).	"Have you been looking for other work in the last 4 weeks?"
National Longitudinal Survey – Young Women (NLSYW)	1985 and 1987	Women aged 33 to 41 that are wage and salary workers with tenure > 1 month (1,883).	"Have you been looking for other work during the past 4 weeks?"
Canada			
Labour Force Survey (LFS)	March 1976-1995	Wage and salary workers with tenure > 1 month (45,372).	"In the past 4 weeks, have you looked for another job?"

Table 2 **On-the-job search rates by age and gender, U.S. and Canada**

	U.S.			Canada		
	1976	1993	Change	1976	1995	Change
Men						
16-19	0.161	0.222	0.061	0.046	0.109	0.063*
20-24	0.082	0.235	0.153	0.037	0.101	0.064*
25-29	0.067	0.136	0.069	0.033	0.078	0.045*
30-39	0.033	0.096	0.063	0.019	0.050	0.031*
40-49	0.021	0.086	0.065	0.010	0.030	0.020*
50 and over	0.012	0.026	0.014	0.005	0.018	0.013*
Women						
16-19	0.083	0.226	0.143	0.047	0.113	0.066*
20-24	0.082	0.158	0.076	0.043	0.112	0.069*
25-29	0.050	0.128	0.078	0.033	0.067	0.034*
30-39	0.042	0.107	0.065	0.013	0.048	0.035*
40-49	0.018	0.113	0.095**	0.012	0.038	0.026*
50 and over	0.012	0.023	0.011	0.005	0.025	0.020*

Note: The U.S. sample is household heads and wives, whereas the Canadian sample is all wage and salary workers.
*,** indicate if changes are statistically significant at the 1 percent and 10 percent level respectively.
Source: May 1976 Current Population Survey and 1993 Panel Study of Income Dynamics for the U.S. March 1976-1995 Labour Force Survey for Canada.

Table 3 **Probit estimates of the probability of on-the-job search**

	1975 PSID[a]		1976 CPS[b]		1976 LFS[c]	
Age 16-19	1.503*	(0.275)	1.388*	(0.130)	0.573*	(0.176)
Age 20-24	1.115*	(0.201)	1.156*	(0.095)	0.637*	(0.172)
Age 25-29	0.819*	(0.200)	0.996*	(0.094)	0.631*	(0.172)
Age 30-39	0.625*	(0.203)	0.774*	(0.093)	0.445*	(0.171)
Age 40-49	0.499*	(0.201)	0.561*	(0.095)	0.389*	(0.176)
Age 50-59	-0.029	(0.216)	0.373*	(0.098)	0.219	(0.186)
Black	0.164	(0.108)	0.018	(0.049)		
Hispanic			-0.110	(0.078)		
Female	0.054	(0.118)	-0.012	(0.061)	-0.013	(0.060)
Married	-0.012	(0.125)	-0.116*	(0.050)	0.074	(0.062)
Female and married			-0.314*	(0.069)	-0.233*	(0.081)
Wife works	0.025	(0.088)			-0.074	(0.055)
Have children	0.033	(0.085)				
High school	0.170	(0.152)				
Some post high school	0.369*	(0.158)				
College degree	0.558*	(0.182)				
High school incomplete			0.114	(0.071)		
High school graduate			0.088	(0.066)		
Post high school			0.301*	(0.070)		
High school					0.071	(0.071)
Some post high school					0.211*	(0.085)
Post-high diploma					0.190*	(0.086)
University degree					0.291*	(0.093)
Reside in house	-0.008	(0.084)	-0.073	(0.057)		
Rent residence	0.074	(0.082)				
Union member	-0.093	(0.084)				
Prefer more hours	0.241*	(0.079)				
Prefer fewer hours	0.150	(0.150)				
Weekly hours 0-14	0.099	(0.318)				
Weekly hours 15-29	0.361*	(0.175)				
Weekly hours 30-40	0.035	(0.074)				
Part-time			0.331*	(0.044)		
Involuntary part-time					1.218*	(0.093)
Voluntary part-time					0.122*	(0.059)
Paid by the hour			0.143*	(0.034)		
Job tenure 2-6 months					1.496*	(0.288)
Job tenure 7-12 months					1.178*	(0.290)

Table 3 - *(continued)* **Probit estimates of the probability of on-the-job search**

	1975 PSID[a]		1976 CPS[b]		1976 LFS[c]	
Job tenure 3-5 years					1.022*	(0.288)
Job tenure 6-10 years					0.810*	(0.290)
Job tenure 11-20 years					0.474	(0.302)
Constant	-1.440*	(0.625)	-2.533*	(0.233)	-4.199*	(0.350)
Pseudo R^2		0.138		0.093		0.142
Number of observations		2,901		33,167		26,797

Note: The regressions also include region/province dummies and industry and occupation dummies.
[a] Sample is employed household heads that are wage and salary workers. Omitted categories are "Northeast," "Age 60 and over," "Primary school," "Weekly hours 40 and over," "Construction industries," and "Farming occupations."
[b] Sample is employed household heads and wives that are wage and salary workers. Omitted categories are "Northeast," "Age 60 and over," "Primary school," "Construction industries," and "Farming occupations."
[c] Sample is employed wage and salary workers. Omitted categories are "Quebec," "Age 60 and over," "Primary school," "Job tenure 20 years and over," "Construction industries," and "Construction occupations."
Standard errors are in parantheses. * indicates signficance at the 5% level.

Table 4 **Decomposition of upward trends in on-the-job search**

A. Overall decomposition

	PSID[a]		CPS/PSID[b]		LFS[c]	
	Actual	Predicted	Actual	Predicted	Actual	Predicted
Early year	0.123	0.123	0.035	0.035	0.023	0.023
Late year	0.240	0.240	0.097	0.097	0.054	0.054
Change		0.117		0.062		0.031
Expected in early year		0.210		0.079		0.045
Due to characteristics		0.030		0.018		0.009
(%)		(26)		(29)		(29)
Due to returns		0.087		0.044		0.022
(%)		(74)		(71)		(71)
Expected in late year		0.136		0.042		0.029
Due to characteristics		0.013		0.007		0.006
(%)		(11)		(11)		(19)
Due to returns		0.104		0.055		0.025
(%)		(89)		(89)		(81)

B. Contribution of each variable to upward trend

	PSID[a]	CPS/PSID[b]	LFS[c]
Province/region	0.001	-0.000	0.001
Age	0.007	0.005	-0.001
Female	-0.001	-0.000	-0.001
Married	0.005	0.004	-0.000
Female and married		-0.001	-0.001
Spouse employed	0.000		0.001
Children	-0.001		
Education	0.009	0.007	0.011
Black	0.001	0.000	
Hispanic		-0.000	
Paid by the hour		0.001	
Reside in house	0.000	0.001	
Rent residence	0.000		
Tenure			-0.007
Usual weekly hours	0.001		
Part-time status		0.000	0.006
Hours preference	0.001		
Union member	0.006		
Industry	-0.001	0.004	0.001
Occupation	0.001	-0.002	-0.000
Total	0.030	0.018	0.009

a Early and late years are 1975 and 1985 respectively. Sample is employed household heads that are wage and salary workers.
b Early and late years are 1976 and 1993 respectively. Sample is employed household heads and wives that are wage and salary workers.
c Early and late years are 1976 and 1995 respectively. Sample is employed wage and salary workers.

Table 5 **On-the-job search rates by tenure, full-time men, Canada**

Job Tenure (years)	1976-1980	1981-1985	1986-1990	1991-1995
0 – 1	0.049	0.058	0.061	0.072
1 – 5	0.022	0.030	0.042	0.043
5 – 10	0.011	0.015	0.024	0.026
10 – 15	0.004	0.006	0.014	0.016
> 15	0.001	0.003	0.006	0.006

Note: The sample is employed male wage and salary workers with usual weekly hours at least 40. All rates are based on samples of at least 10,000 observations.
Source: March 1976-1995 Labour Force Survey.

Table 6 **Within-cohort on-the-job search rates, United States**

Year of birth	1976-1977	1988-1993	Change
Before 1940	0.018	0.029	0.011
1940 – 1949	0.047	0.072	0.025
1950 – 1959	0.079	0.093	0.015
After 1959	0.211	0.141	-0.070*

Note: * indicates if change in rates is statistically significant at the 5 percent level.
Source: May 1976 and 1997 Current Population Survey and 1988-1993 Panel Study of Income Dynamics. The sample is employed household heads and wives that are wage and salary workers.

Table 7 **Probit estimates of the probability of an employer change**

	(1)	(2)	(3)
Year 1988 dummy	0.2187* (0.0415)	0.2671* (0.0594)	0.1984* (0.0614)
Weeks since last interview		0.0150* (0.0040)	0.0151* (0.0041)
Local unemployment rate		-0.0000 (0.0132)	0.0035 (0.0134)
Age			-0.0685* (0.0093)
Married			-0.2209* (0.0451)
Black			0.0305 (0.0823)
Some or completed college			-0.0851 (0.0439)
Enrolled in school			0.1205 (0.0736)
No. of observations	4,405	4,363	4,363

Note: Standard errors are in parentheses. * indicates significance at the 5% level.
Source: Sample of employed men aged 23-31 that are wage and salary workers in their main jobs in 1975 NLSYM and 1988 NLSY.

Table 8 **On-the-job search methods, Canada, 1976-1981**

	1976	1977	1978	1979	1980	1981
Contacted employer	0.016	0.018	0.020	0.023	0.023	0.024
Contacted public agency	0.007	0.008	0.008	0.008	0.008	0.008
Looked at ads	0.006	0.007	0.008	0.009	0.009	0.010
Other	0.010	0.011	0.010	0.011	0.011	0.012
On-the-job search rate	0.026	0.028	0.030	0.034	0.034	0.036
Average number of methods	1.54	1.57	1.56	1.50	1.52	1.50

Note: Sample is *all* employed workers, including the self-employed.
Source: The Canadian data are from Statistics Canada, *Labour Force Information*, Catalogue no. 71-001, monthly 1976-1995.

Figure 1
On-the-job search rates, U.S., 1969-1995

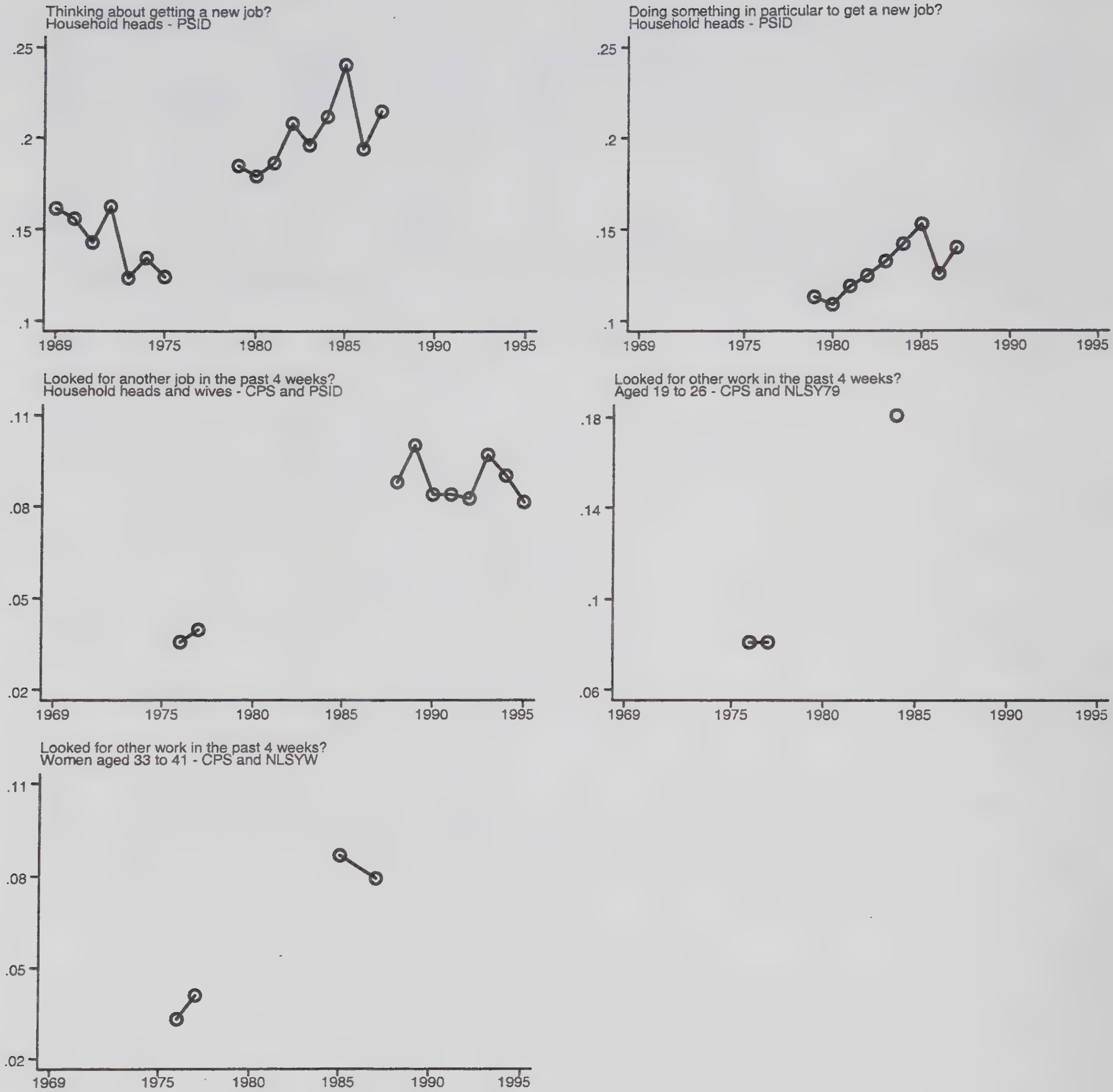


Figure 2
On-the-job search rates, Canada, 1976-1995

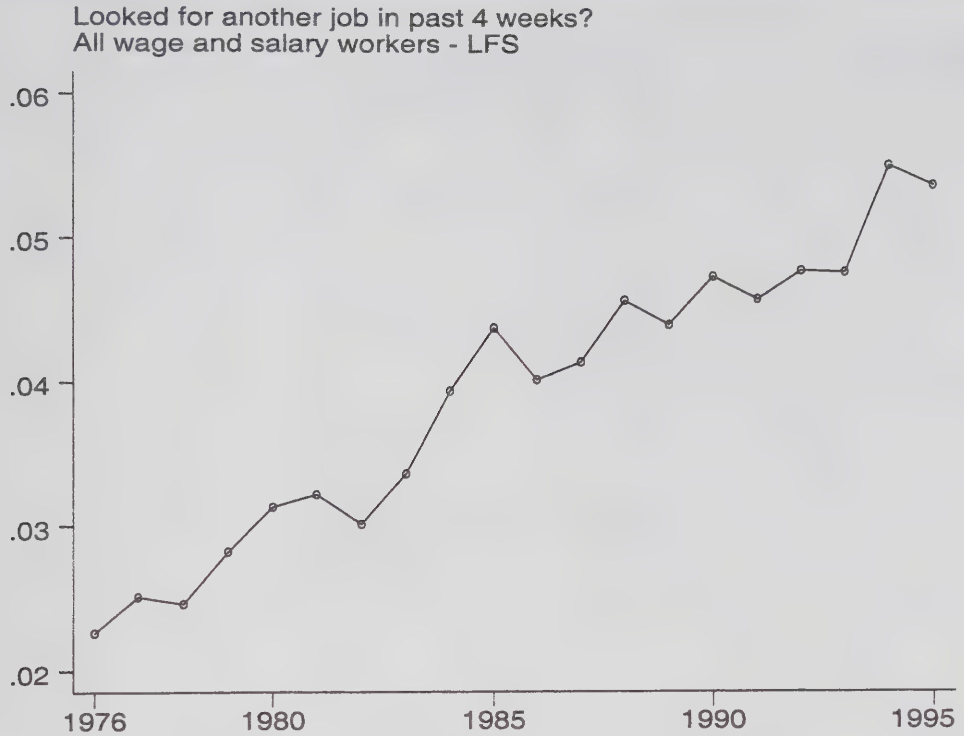


Figure 3
Within-cohort on-the-job search rates, Canada

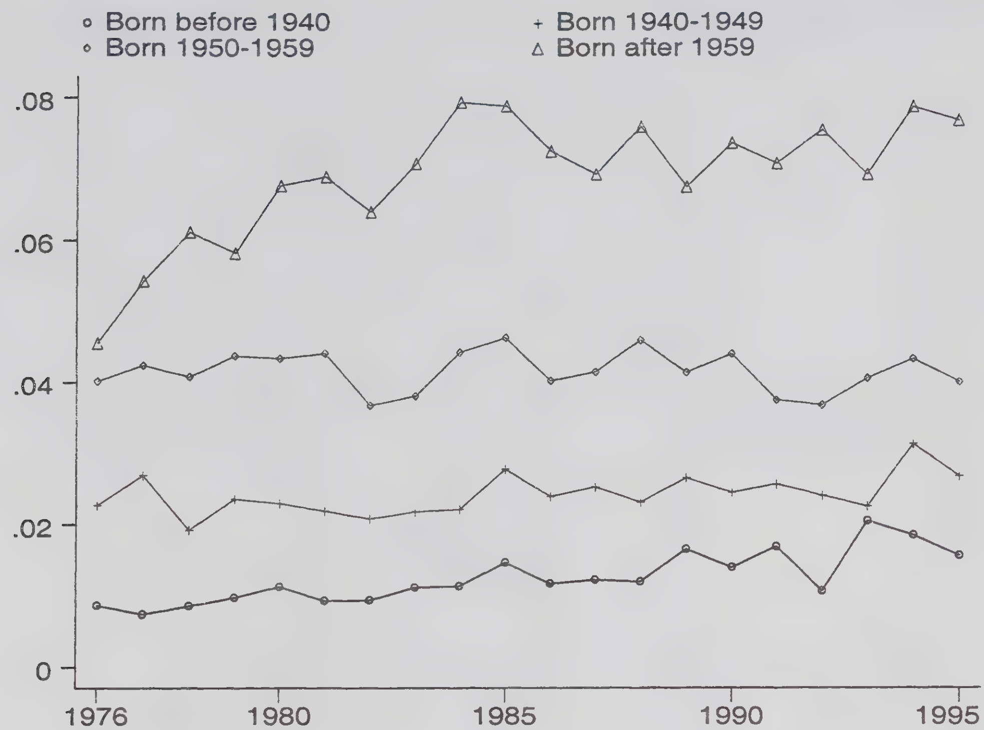
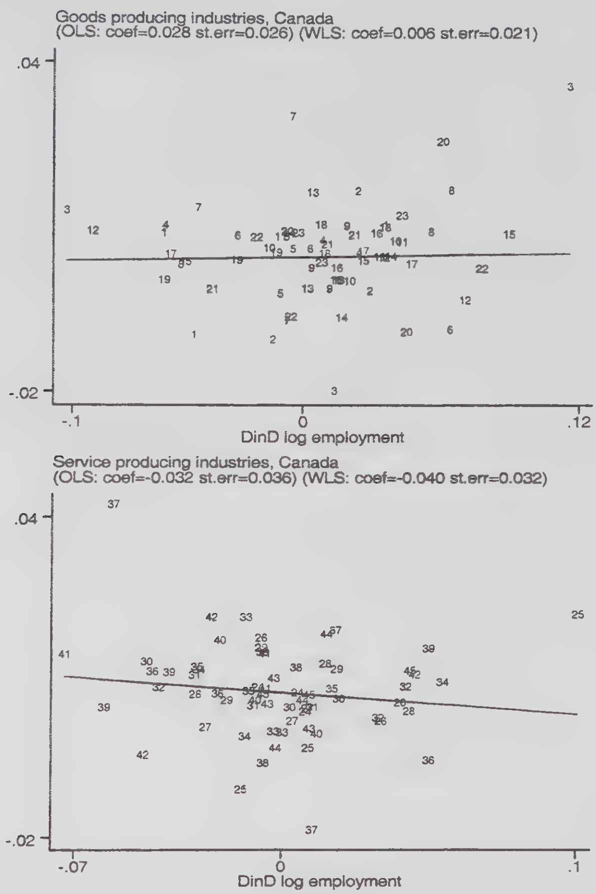
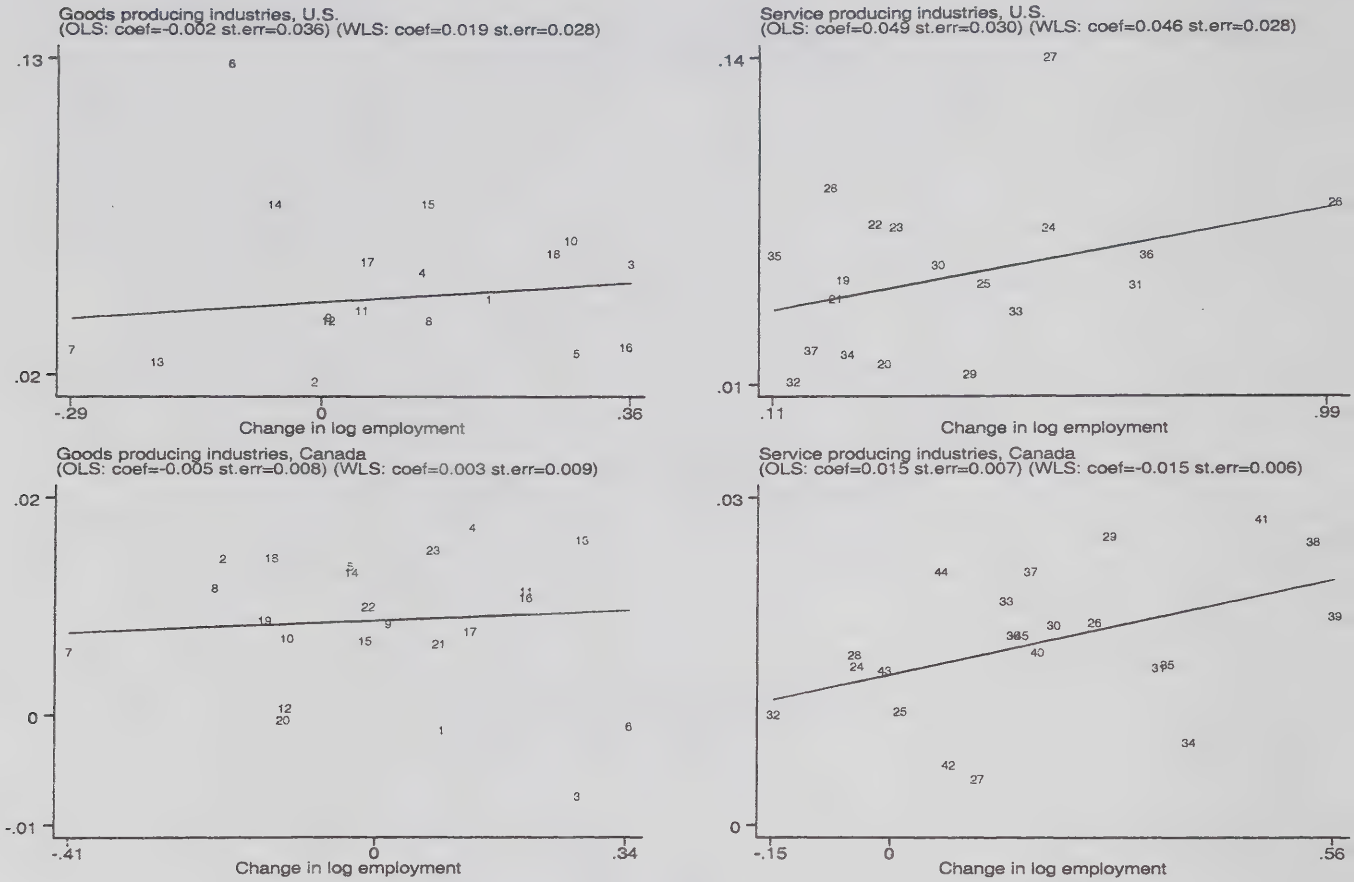


Figure 4
Changes in on-the-job search rates and differences-in-differences log employment



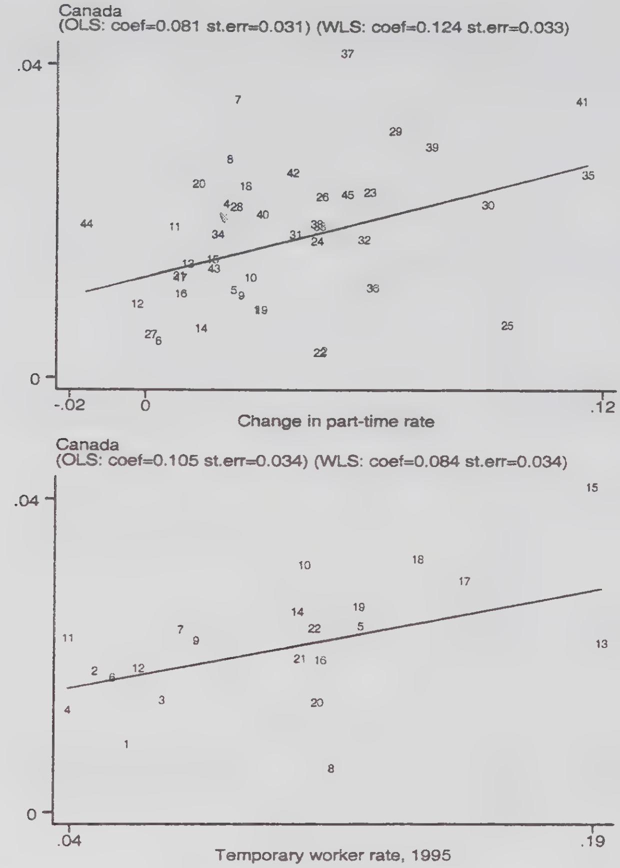
Note: Vertical axes measure percentage point changes in the OJS rate. The regression line is from the WLS estimates. Industry codes are in Table A1. Differences-in-differences log employment is the difference in the mean annual absolute log employment change between two five-year periods. Data on four five-year periods are plotted: 1976-1980, 1981-1985, 1986-1990 and 1991-1995.
Source: Change in Canadian OJS rates and log employment levels are from March 1976-1995 Canadian Labour Force Surveys.

Figure 5
Changes in on-the-job search rates and log employment by industry, mid-1970s to late-1980s



Note: Industry codes for the Canada and the U.S. are in Tables A1 and A2 respectively. Vertical axes measure percentage point changes in the OJS rate. The regression line is from the WLS estimates.
Source: Change in U.S. OJS rates are from May 1976 and 1977 Current Population Survey and 1988-1993 Panel Study of Income Dynamics. Change in U.S. log employment levels is from the 1976 and 1988 basic monthly Current Population Survey files. Change in Canadian OJS rates and log employment levels are from March 1976-1980 and 1986-1990 March Labour Force Surveys.

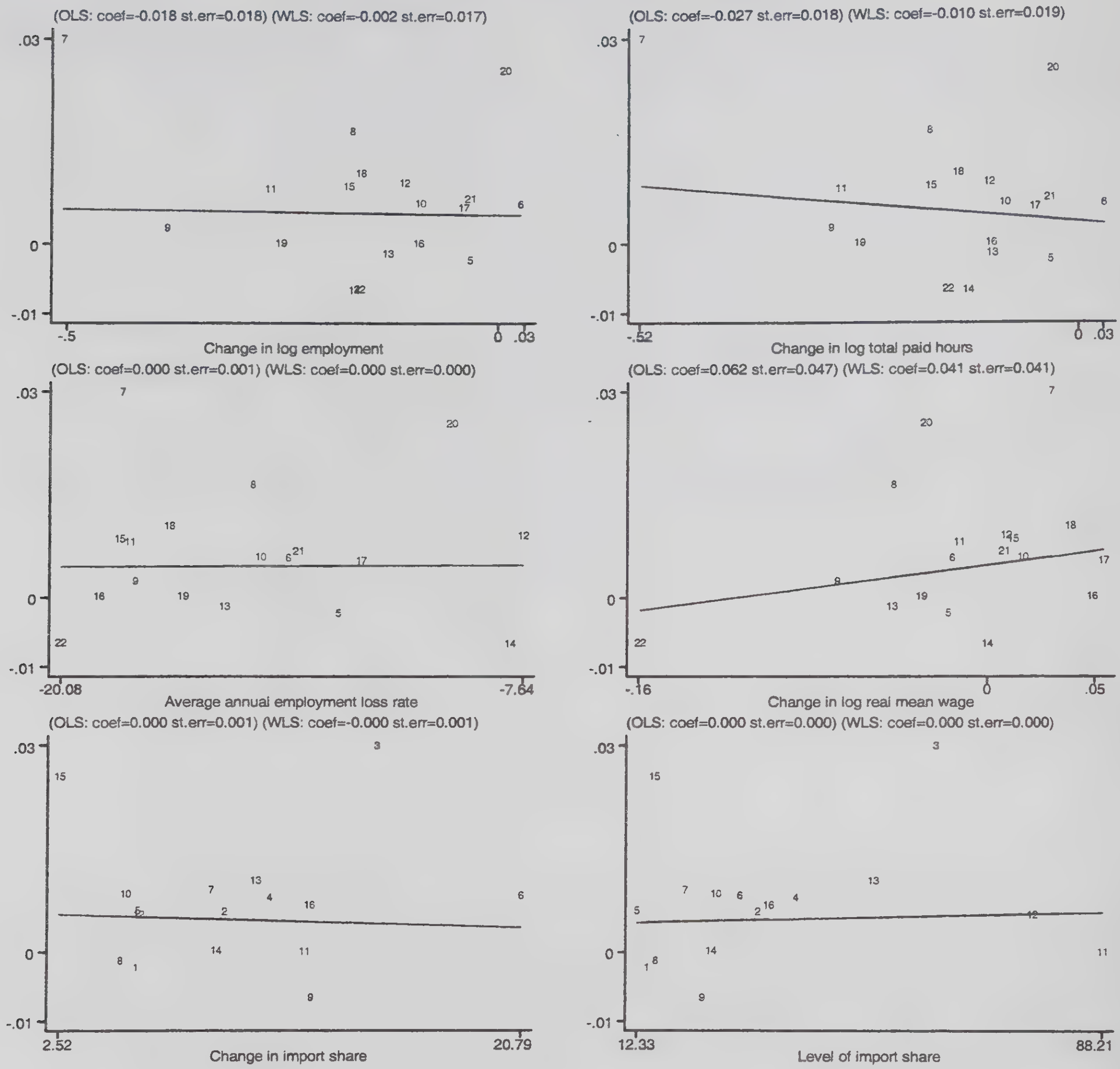
Figure 6
Changes in on-the-job search rates and part-time/contingent worker rates



Note: Vertical axes measure percentage point changes in the OJS rate. The regression line is from the WLS estimates. Industry codes for panel 1 are in Table A1. Industry 3 (Fishing and trapping) is an outlier in panel 1 and is dropped from the graph (change in part-time rate = 0.239 and change in OJS rate = 0.028). Industry codes for panel 2 are: 1. Agriculture 2. Primary industries 3. Non-durable manufacturing 4. Durable manufacturing 5. Construction 6. Transportation and storage 7. Communications 8. Utilities 9. Wholesale trade 10. Retail trade 11. Finance industries 12. Insurance and real estate 13. Education and related services 14. Health and welfare services 15. Amusement and recreation 16. Services to business management 17. Personal services 18. Accommodation and food services 19. Miscellaneous services 20. Federal administration 21. Provincial administration 22. Local administration.
Source: Change in OJS and part-time rates are from March 1976-1980 and 1991-1995 March Labour Force Surveys. Temporary worker rate is from the Survey of Work Arrangements supplement to the November 1995 Labour Force Survey.

Figure 7

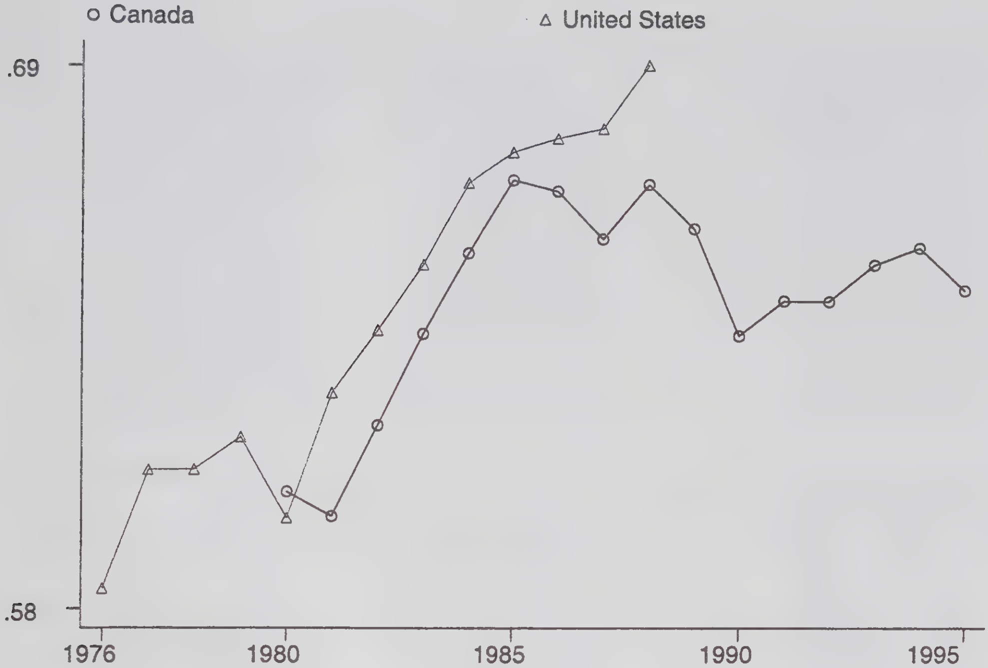
Changes in on-the-job search rates and production worker employment/wages and industry trade, late-1980s to early 1990s, Canadian manufacturing industries



Note: Vertical axes measure percentage point changes in the OJS rate. The regression line is from the WLS estimates. Industry codes for the first four panels are in Table A1. Industry codes for the fifth and sixth panel are: 1. Food, beverage, tobacco 2. Rubber and plastic 3. Leather 4. Textiles and clothing 5. Wood 6. Furniture and fixtures 7. Paper and allied 8. Printing and publishing 9. Primary metal 10. Fabricated metal 11. Machinery 12. Transportation equipment 13. Electrical machinery 14. Non-metallic mineral products 15. Petroleum and coal 16. Chemical.

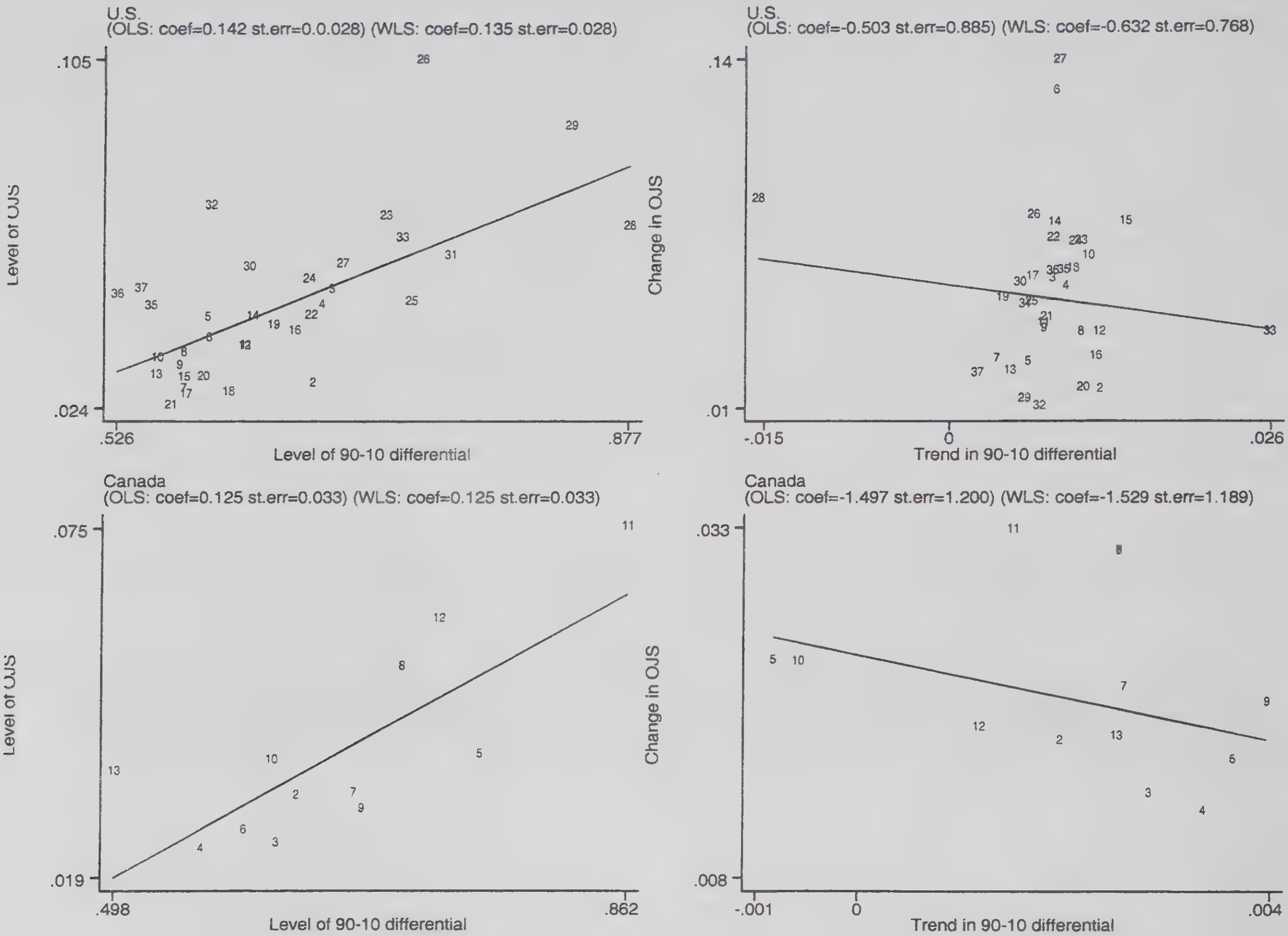
Source: Change in Canadian OJS from March 1986-1990 and 1991-1995 March Labour Force Surveys. Employment and wage changes in panels 1-4 from Annual Survey of Manufactures 1986-1990 and 1991-1995. Employment loss rate in panel 4 from Annual Survey of Manufactures 1976-1995. This series is calculated as the loss in employment from all plant contractions and closings between two years as a function of total employment in the first year. Trade data are from are from the System of National Accounts (Statistics Canada, *The Input-Output Structure of the Canadian Economy*, 1986-1993).

Figure 8
90-10 differentials of absolute within-industry log wage residuals



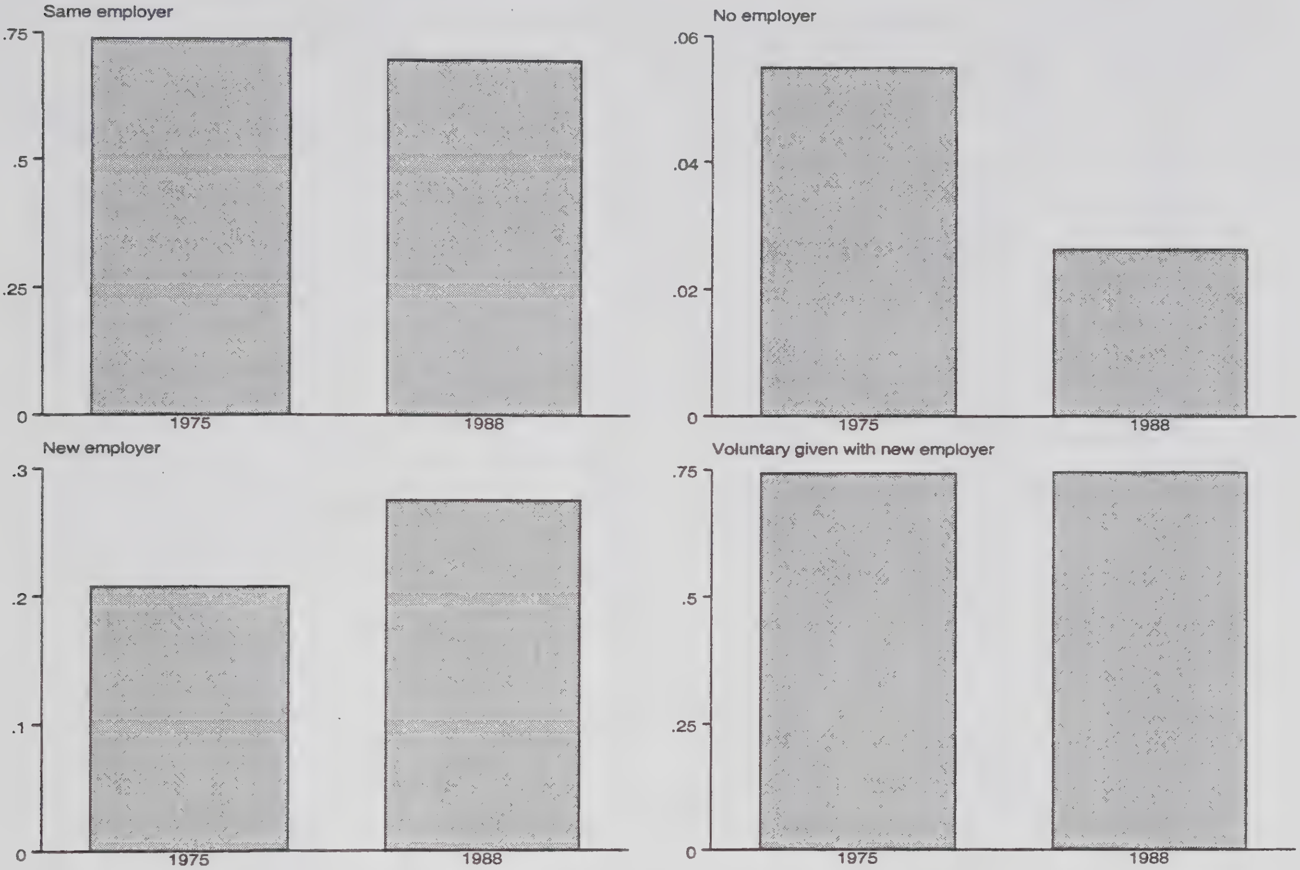
Note: National differentials are created by taking a weighted average of the industry differentials, where the weights are the industry employment shares in each year.
Source: March Current Population Survey 1977-1989 for the U.S. and Survey of Consumer Finances 1981-1996 for Canada.

Figure 9
Level of and change in on-the-job search rates and levels and trends in 90-10 differentials of absolute log wage residuals, non-agricultural industries



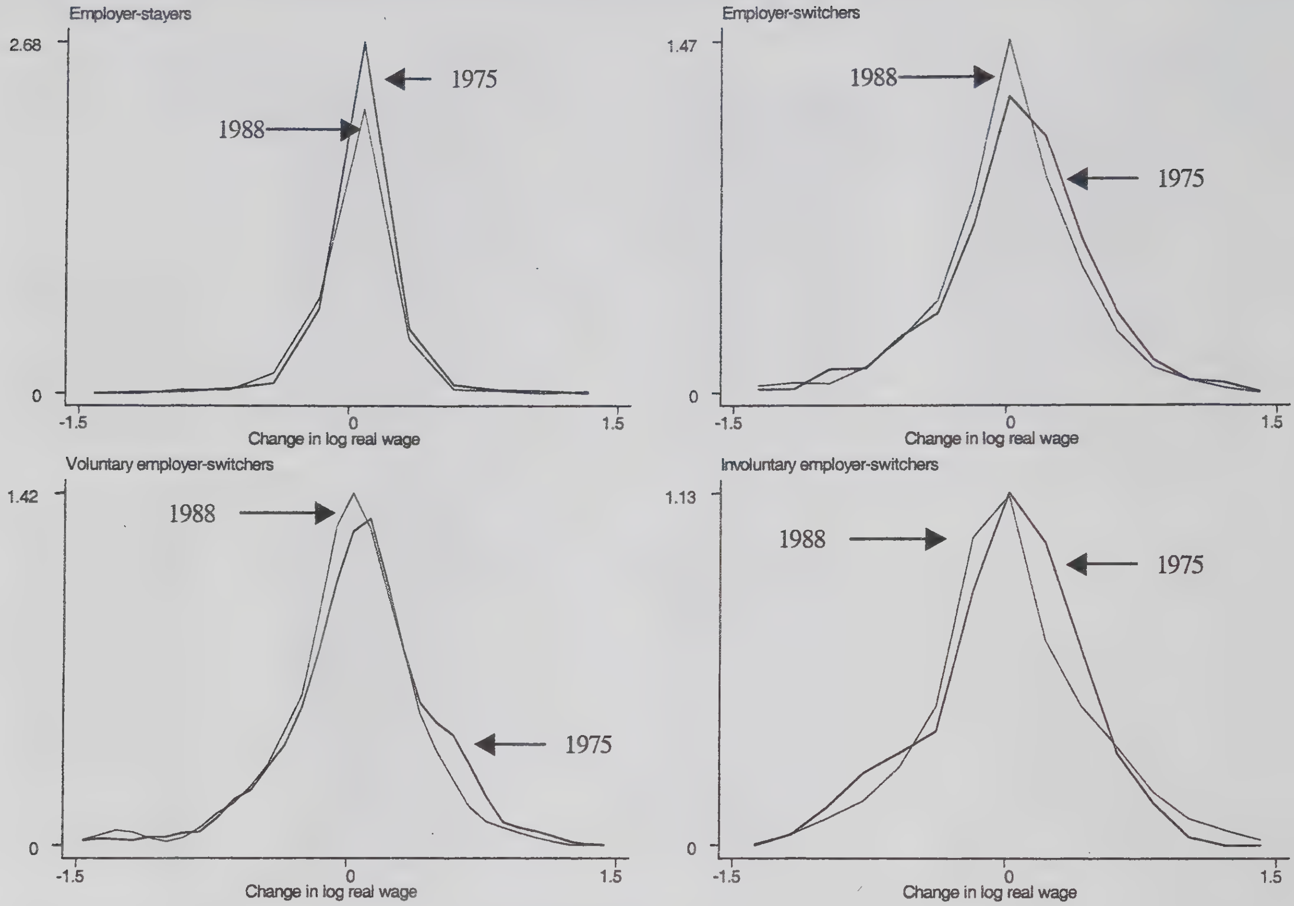
Note: Industry codes for the U.S. are in Table 8. Separate wage data for industry 34 (postal service) is not available. Industry codes for Canada: 1. Agriculture 2. Primary industries 3. Non-durable manufacturing 4. Durable manufacturing 5. Construction 6. Transportation, communications and other utilities 7. Wholesale trade 8. Retail trade 9. Finance, insurance and real estate 10. Community services 11. Personal services 12. Business and miscellaneous services 13. Public administration. The regression line is from the WLS estimates.
Source: Change in U.S. OJS rates are from May 1976 and 1977 Current Population Survey and 1988-1993 Panel Study of Income Dynamics. U.S. wage dispersion data from March Current Population Survey 1977-1989. Change in Canadian OJS rates are from March 1976-1980 and 1991-1995 March Labour Force Surveys. Canadian wage dispersion measures are from the Survey of Consumer Finances 1981-1996.

Figure 10
One-year transition rates, employed male wage and salary workers aged 23-31



Source: Sample of employed men aged 23-31 that are wage and salary workers in their main jobs in 1975 NLSYM and 1988 NLSY

Figure 11
Distributions of one-year real log wage changes



Note: The distributions are estimated using a Epanechnikov kernel density function. The bandwidth of the kernel is calculated as $h = 0.9m / n^{1/5}$ where n is the number of observations and m is the standard deviation of x.

Appendix A **Period-to-period changes in OJS rates, by industry**

Table A1 **On-the-job search rates by industry, Canada.**

	1976-1980	1981-1985	1986-1990	1991-1995
Goods-producing industries				
1. Agriculture	0.033	0.042	0.032	0.042
2. Forestry	0.034	0.051	0.049	0.038
3. Fishing and trapping	0.025	0.038	0.018	0.053
4. Mining, quarries and oil wells	0.020	0.027	0.037	0.042
5. Food, beverage, tobacco	0.019	0.024	0.032	0.030
6. Rubber and plastics	0.026	0.016	0.024	0.030
7. Leather	0.016	0.029	0.022	0.051
8. Textiles	0.010	0.018	0.021	0.037
9. Clothing	0.014	0.012	0.022	0.024
10. Wood	0.027	0.026	0.033	0.039
11. Furniture and fixtures	0.019	0.023	0.030	0.038
12. Paper and allied	0.009	0.013	0.010	0.018
13. Printing, publishing and allied	0.037	0.037	0.052	0.051
14. Primary metal	0.009	0.013	0.022	0.015
15. Fabricated metal	0.022	0.026	0.029	0.037
16. Machinery	0.020	0.022	0.030	0.030
17. Transportation equipment	0.011	0.014	0.019	0.024
18. Electrical products	0.019	0.023	0.033	0.043
19. Non-metallic mineral products	0.020	0.025	0.029	0.029
20. Petroleum and coal	0.009	0.017	0.008	0.033
21. Chemical	0.015	0.023	0.021	0.027
22. Miscellaneous	0.018	0.026	0.028	0.021
23. Construction industries	0.027	0.039	0.042	0.051
Service-producing industries				
24. Transportation	0.021	0.028	0.036	0.039
25. Storage	0.018	0.039	0.028	0.024
26. Communications	0.013	0.015	0.032	0.036
27. Electric power, gas and water utilities	0.014	0.014	0.018	0.019
28. Wholesale trade	0.022	0.034	0.038	0.044
29. Retail trade	0.035	0.050	0.061	0.066
30. Finance	0.013	0.026	0.031	0.035
31. Insurance and real estate	0.024	0.033	0.038	0.042
32. Elementary and secondary schools	0.018	0.020	0.028	0.036
33. Universities and colleges	0.043	0.064	0.064	0.063
34. Other education	0.044	0.052	0.051	0.062
35. Health and welfare services	0.023	0.030	0.038	0.049
36. Religious organizations	0.013	0.024	0.030	0.024
37. Amusement and recreation services	0.058	0.099	0.080	0.099
38. Services to business management	0.041	0.056	0.067	0.060
39. Personal services	0.027	0.042	0.046	0.056
40. Accommodation services	0.059	0.075	0.074	0.080
41. Food services	0.060	0.074	0.088	0.095
42. Miscellaneous services	0.063	0.058	0.068	0.089
43. Federal administration	0.033	0.037	0.047	0.047
44. Provincial administration	0.025	0.030	0.048	0.044
45. Local and other government services	0.018	0.025	0.035	0.042

Source: March 1976-1995 Labour Force Survey. The sample is employed wage and salary workers.

Table A2 **On-the-job search rates by industry, U.S.**

	1976-1977	1988-1993
Goods-producing industries		
1. Agriculture	0.039	0.084
2. Other primary	0.027	0.043
3. Construction	0.038	0.096
4. Lumber and wood products	0.035	0.090
5. Furniture and fixtures	0.035	0.061
6. Stone, clay and glass	0.025	0.153
7. Metal industries	0.023	0.051
8. Machinery	0.028	0.066
9. Electrical machinery, equipment and supplies	0.024	0.063
10. Transportation	0.021	0.087
11. Miscellaneous manufacturing	0.030	0.072
12. Food, beverage, tobacco	0.030	0.068
13. Textile mill products	0.030	0.054
14. Apparel	0.024	0.102
15. Paper and allied products	0.019	0.098
16. Printing and publishing	0.028	0.057
17. Chemical, Petroleum and coal products	0.018	0.076
18. Rubber and plastics	0.017	0.078
Service-producing industries		
19. Transportation services	0.030	0.081
20. Communications	0.024	0.041
21. Utilities and sanitary services	0.015	0.058
22. Wholesale trade	0.029	0.102
23. Retail trade	0.047	0.118
24. Finance	0.032	0.103
25. Insurance and real estate	0.040	0.088
26. Business services	0.065	0.146
27. Repair services	0.022	0.161
28. Personal services	0.045	0.132
29. Entertainment and recreation	0.088	0.101
30. Health services	0.035	0.091
31. Welfare and religious services	0.040	0.088
32. Education	0.065	0.075
33. Other professional services	0.054	0.092
34. Postal service	0.011	0.031
35. Federal administration	0.027	0.087
36. State administration	0.032	0.093
37. Local administration	0.041	0.063

Source: May 1976 and 1997 Current Population Survey and 1988-1993 Panel Study of Income Dynamics. The sample is employed household heads and wives that are wage and salary workers.

Appendix B **Weighted least squares estimator**

The weighted least squares (WLS) estimates supplement the OLS estimates by weighting observations based on larger cell sizes more heavily. We are interested in estimating the equation:

$$\Delta p_i = \alpha + \beta x_i + v_i, \tag{1}$$

where Δp_i is the true change in OJS rates experienced by industry i, x_i is the magnitude of some other change experienced by industry i over the same period, and v_i is a random error with expected value 0 and a uniform variance σ^2. The problem is, of course, that we do not observe Δp_i. Instead, we must estimate:

$$\Delta \hat{p}_i = \alpha + \beta x_i + e_i \tag{2}$$

where $\Delta \hat{p}_i$ is the estimated change in OJS rates experienced by industry i, and β are the OLS estimates presented in Figures 4, 5, 6, 8 and 9. The error term in (2), e_i, does not have a uniform variance. Rather,

$$Var(e_i) = Var(p_{il}) + Var(p_{ie}) + \sigma^2 \tag{3}$$

where p_{il} and p_{ie} are the OJS rates for industry i from the early and late periods respectively. To correct for this heteroscedasticity, the estimate of β is instead obtained from:

$$\frac{\Delta \hat{p}_i}{\sqrt{w_i}} = \frac{\alpha}{\sqrt{w_i}} + \frac{\beta x_i}{\sqrt{w_i}} + \frac{e_i}{\sqrt{w_i}} \tag{4}$$

where w_i is given by the right-hand-side of (3) and an estimate of σ^2 is obtained from:

$$\hat{\sigma}^2 = \frac{\sum \hat{e}_i^2}{n-1} - \frac{\sum (Var(\hat{p}_{il}) + Var(\hat{p}_{ie}))}{n}. \tag{5}$$

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